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Are Speculative Attacks Triggered by Sunspots? A New Test*

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Abstract

The empirical methodology of the paper establishes if a speculative attack, which is accounted for via sunspots in the presence of multiple equilibria, could have been in fact driven uniquely by economic fundamentals. The methodology is based on the theoretical models of Bertola and Svensson (1993) and Tarashev (2003). The first model captures robust stylised facts from target zone regimes, whereas the second one implies that both unique and multiple equilibria can account for violent speculative attacks. The characteristics of the theoretical foundations and their implications for the employed statistical test distinguish the paper from previous structural empirical analyses of market bets against pegged currencies. The methodology is applied to the experience of two ERM countries in the fall of 1992. The attack on the French Franc is found to be triggered by sunspots, whereas it is impossible to determine whether a similar scenario or the state of the economy alone underpins the currency crisis in Italy.

JEL Classification Numbers: C22, D84, F31

*The views expressed herein are those of the author and do not necessarily reflect those of the BIS.

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1 Introduction

When market players bet against a pegged currency, they re-evaluate their investment positions and engage in what is commonly known as a speculative attack. The attack puts pressure on exchange and interest rates and, if successful, leads to a currency crisis: a collapse of the exchange rate regime.

The 1990s witnessed a series of speculative attacks, some of which were characterised as violent because their intensity seemed out of line with the concurrent behaviour of economic fundamentals. In practical terms, an attack is viewed as violent if its intensity, measured by the induced changes in interest and exchange rates, cannot be explained well by a linear function of the underlying fundamentals, which include measures of the country's competitiveness, unemployment rate and/or policy stance. There is a general agreement on the timing of such attacks, some of which took place in the run up to the ERM crisis in 1992-3 and the Asian crisis in 1997.[1]

Violent speculative attacks have been the focus of a vast literature because, successful or not in causing currency crises, they disrupt economic activity and lead to substantial losses for market participants.[2] Research in international finance developed the first- and second-generation approaches to currency crises, pioneered respectively by Krugman (1979) and Obstfeld (1994). These theoretical models demonstrate that, before an attack materialises, rational behaviour might reflect only to a limited extent traders' awareness of a deterioration in the fundamentals. At the time of the attack, however, traders' actions fully match the precarious state of the currency regime. The upshot is a phenomenon that qualifies a speculative attack as violent: an apparent decoupling of market outcomes from fundamentals.

Despite the existence of well established theoretical explanations of violent speculative attacks, the theory-based empirical analysis of such attacks is still at an early stage. This motivates the methodology that I develop below and then apply, on an episode-by-episode basis, to the 1992 attacks on the Italian Lira and the French Franc. The empirical procedure is based on two theoretical models that complement each other. One of the models is from Tarashev (2003) and builds on Morris and Shin (1999), which incorporates heterogeneous beliefs in the traditional second-generation approach.[3] That

[1] See, for example, Eichengreen, Rose and Wyplosz (1996). ERM stands for the Exchange Rate Mechanism in Europe.

[2] For a detailed discussion of the issue, refer to Eichengreen (2000), Obstfeld (1994, 1996) or Obstfeld and Rogoff (1995).

[3] I choose not to consider the first-generation approach for two reasons. First, some of its key empirical implications seem to be at odds with the data. Such a conclusion could be drawn from the results of Rose and Svensson (1994) Kaminsky, Lizondo and Reinhart (1998) and Kaminsky (1999). Second, the approach postulates that the currency regime is unsustainable in the long run: an assumption that is

model determines the link between fundamentals and endogenous devaluation expectations, which constitute a measure of the intensity of a speculative attack. The second model is developed in Bertola and Svensson (1993) and pins down the exchange rate in a target-zone regime: such were the regimes in Italy and France in the early 1990s when the Lira and the Franc were allowed to fluctuate in currency specific bands around a central parity set vis-à-vis the Deutsche Mark. The latter model captures stylised "target zone" facts regarding the univariate and joint distributions of interest and exchange rates.[4]

The empirical analysis attempts to distinguish between different types of violent speculative attacks by conducting a hypothesis test regarding the attacks' underlying equilibrium. The specification of the test is determined by the model of Tarashev (2003), according to which a violent speculative attack could be driven solely by economic fundamentals (a unique equilibrium outcome) or be the result of self-fulfilling prophecies (a particular outcome out of several possible equilibria). The null hypothesis of the test postulates equilibrium uniqueness, which implies that the fundamentals have to deteriorate to a threshold value in order to trigger an attack. Owing to the monotone relationship between fundamentals and devaluation expectations, the null is rejected if and only if there is insufficient evidence that devaluation expectations reach their highest pre-attack level just before the attack. Failure to reject the null leaves the test inconclusive. Multiplicity of equilibria is shown to be an irrefutable hypothesis because it implies that an attack could be triggered by economically meaningless and unidentifiable sunspots for any behaviour of the fundamentals.

The type of equilibrium underlying market behaviour has important policy implications. Multiplicity implies that a central authority could avoid a crisis by coordinating market participants' actions on the "better" outcome. Uniqueness, on the other hand, suggests that crisis management should focus exclusively on fundamentals.

To my knowledge, all of the existing theory-based tests of speculative attacks incorporate the empirical procedure of Jeanne (1997) or Jeanne and Masson (2000) and thus hinge on the predictions of the traditional representative-agent version of the second-generation approach to currency crises. Such tests are inherently biased because the traditional version of the approach captures violent speculative attacks only under multiple equilibria. This stands in contrast to the empirical exercise in this paper and casts doubt on the conclusions of existing theory-based analyses which infallibly find that

probably too strong in the context of the ERM in 1992-3. Obstfeld and Rogoff (1995) and Eichengreen (2000) argue that speculative attacks on the ERM targeted currecy regimes whose collapse was not inevitable.

[4] These stylised facts are spelt out in Sections 3 and 4 and were first reported by Bertola and Svensson (1993) and Garber and Svensson (1995).

violent speculative attacks are driven by sunspots.[5]

In the context of the empirical literature on speculative attacks, the test methodology developed in this paper has new implications for the employed data set. The null hypothesis, expressed exclusively in terms of the behaviour of devaluation expectations prior to an attack, demands only the use of pre-attack interest and exchange rates. Available at the daily frequency, these variables would typically provide richer information about rapidly evolving foreign exchange markets than the traditionally used macroeconomic fundamentals, which are observed at monthly or lower frequencies.

The theoretical underpinnings of the empirical exercise impose certain characteristics on the employed data. The exclusion of time periods, which are characterised by economic phenomena unaccounted for by the models, limits the sample size. In addition, the framework of Tarashev (2003) makes it imperative to consider financial contracts with non-overlapping payoff horizons. Since the observations are daily, this translates into a requirement to use overnight interest rates.

The model-imposed features of the data affect the empirical procedure. To estimate devaluation expectations, one needs an estimate of the speed of mean reversion in the underlying stochastic variables. The mean reversion parameters are, however, estimated with substantial uncertainty: these parameters relate to low-frequency processes, the information about which is limited by the sample size and the horizon of the interest rates. To address the issue, standard inference procedures need to be replaced with the ones developed by Stock (1991).

The data suggest different interpretations for the two episodes under study. Market expectations of a devaluation of the Italian Lira increase in a sustained fashion during the last three quarters of the sample and settle at their maximum level just before the attack. The attack is thus consistent with both uniqueness and multiplicity of equilibria. In contrast, the data provide no evidence that the expectations of a devaluation of the French Franc increase in the run-up to the attack on that currency. Consequently, the statistical test prompts a rejection of the hypothesis that the episode was the outcome of a unique equilibrium.

The paper is organized as follows. Section 2 reviews briefly extant theory-based tests of speculative attacks and then specifies the null hypothesis of the test developed in this paper. Section 3 develops the underlying theoretical model. Assuming equilibrium uniqueness over the entire state space, I solve and interpret the model in Section 4. The data are described and motivated in Section 5. Section 6 motivates while Section 7 outlines the empirical procedure. Section 8 reports and interprets the empirical results.

[5] In addition to Jeanne (1997) and Jeanne and Masson (2000), see Boinet, Napolitano and Spagnolo (2003) and Ratti and Seo (2003). I revisit the issue in the next section.

Finally, Section 9 argues rigorously why it is impossible to refute the hypothesis that multiple equilibria drive a violent speculative attack on a target zone regime.

2 Equilibrium uniqueness versus multiplicity in structural tests. Specifying the null

I underscore the innovative features of the test conducted herein by specifying its null hypothesis after discussing related empirical procedures from the literature. Extant studies have invariably relied on the second-generation approach to currency crises in order to provide theory-based accounts of whether speculative attacks are driven only by economic fundamentals or are the outcome of multiple equilibria.

In its traditional form, the approach models a representative private speculator and a central authority administering the exchange rate regime. The beliefs of that speculator could be self-fulfilling because an action motivated by high (low) devaluation expectations influences the authority's objectives in such a way as to increase (decrease) the actual likelihood of a devaluation. The self-fulfilling beliefs decouple market developments from the fundamentals by giving rise to multiple equilibria, each one of which can occur for the same state of the economy, at the whim of economically meaningless sunspots.

Even though it accounts for violent speculative attacks under multiple equilibria, the traditional version of the second-generation approach is not in a position to explain such episodes solely on the basis of economic fundamentals. Figure 1, in which devaluation expectations measure the intensity of an attack, provides an illustration of the equilibrium implications of that version of the approach.[6] The smooth function in the left panel of the figure implies that, in the absence of a jump in the fundamentals, an attack could evolve only gradually when the equilibrium is unique.[7] The implications of multiple equilibria are quite different. As long as the fundamentals are in the interval delimited by the dashed lines in the right panel, a sunspot would generate an abrupt change in devaluation expectations (and, thus, in market behaviour) by shifting the equilibrium from the bottom to the top branch of the schedule.

Jeanne (1997) and Jeanne and Masson (2000) predicate their tests of violent speculative attacks on the theoretical implications illustrated in Figure 1. As such, the tests are biased towards rejecting the hypothesis of equilibrium uniqueness: not surprisingly,

[6] Without loss of generality, I consider only positive devaluation expectations.

[7] Throughout the paper I implicitly rule out the possibility that a large shock to y could trigger a sudden speculative attack. As argued by Obstfeld (1996), and Eichengreen (2000), there seems to be no evidence for such a shock during the attacks that the paper focuses on.

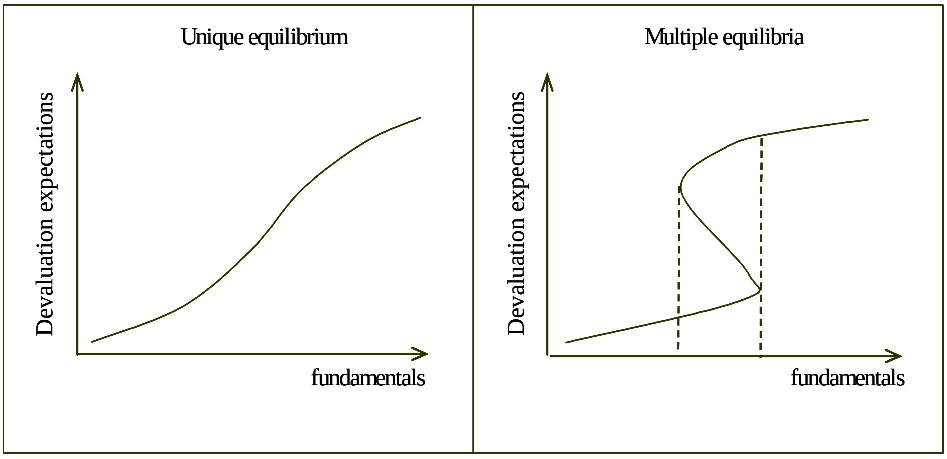

Figure 1: Speculative attacks in the traditional second-generation approach

the two papers conclude that the 1992 attack on the French Franc is driven by multiple equilibria. The same criticism pertains to the multiple-equilibria verdict reached by Boinet, Napolitano and Spagnolo (2003), who apply the test of Jeanne and Masson (2000) to the Argentine crisis in 2002, and by Ratti and Seo (2003), who analyse the 1997 speculative attack in Korea via the procedure of Jeanne (1997).

In this paper, I construct a test of violent speculative attacks that incorporates the theoretical predictions of Tarashev (2003). Let us denote by y the macroeconomic fundamentals that could influence the authority in its decision to abandon or not the current exchange rate regime. If devaluation expectations on date t are denoted by g_t, the model of Tarashev (2003) implies

$$g_t = g\left(y_t\right) \tag{1}$$

where $g\left(\cdot\right)$ represents a function if the equilibrium is always unique but a correspondence if there are multiple equilibria.

The properties of $g\left(\cdot\right)$, illustrated in Figure 2, reflect implications of the model in Morris and Shin (1999), which generalises the traditional second-generation approach by allowing private speculators to hold different beliefs about economic fundamentals. The bigger the importance of private beliefs in the decision making process, the more difficult it is for speculators to coordinate on different equilibria by the means of a sunspot. This is at the root of a strongly non-linear functional relationship between fundamentals and devaluation expectations, which stands in sharp contrast to the implications of the traditional second-generation approach and is illustrated in the left panel of Figure 2.

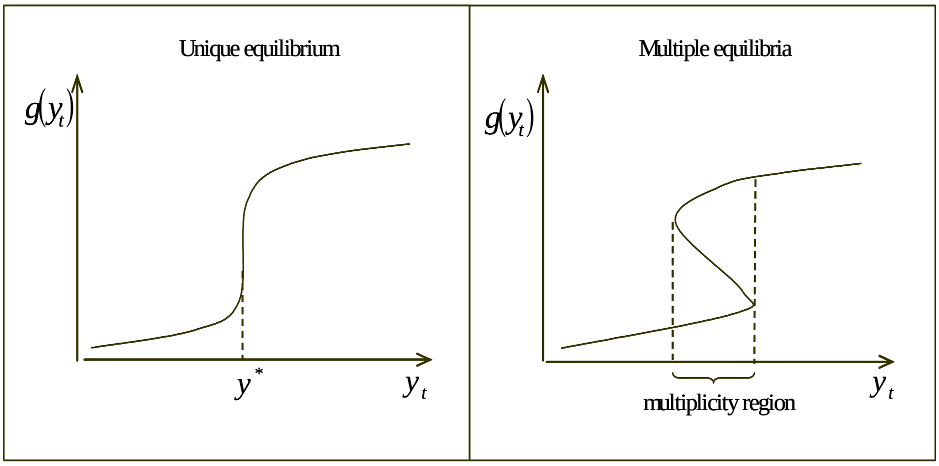

Figure 2: Speculative attacks in Tarashev (2003)

In the unique-equilibrium scenario, a small shock to y_t could tip the economy to the right of y^* where devaluation expectations explode and mark the beginning of a violent speculative attack. These implications are observationally similar to the decoupling of fundamentals from market behaviour under multiple equilibria: the scenario illustrated in the right panel of the figure and relevant when private beliefs play a small role in the decision making process of speculators.

Tarashev (2003) demonstrates that the implications of Morris and Shin (1999) hold true in a context in which devaluation expectations are expressed publicly via market prices. The two scenarios illustrated in Figure 2 could thus be nested in the empirical model of Bertola and Svensson (1993), which allows for evaluating the intensity of a speculative attack by extracting devaluation expectations from observable interest and exchange rates. The nesting, which preserves the implications of the two constituent models, and its underlying assumptions are described in Appendix 1. The resulting setup is outlined in Section 3 and forms the basis of the empirical analysis.

The paper develops a statistical test whose null hypothesis postulates equilibrium uniqueness. The left panel of Figure 2 suggests a way of translating the null into a requirement that can be applied to the data: *devaluation expectations are to reach their highest pre-attack level just before the attack*, i.e., just before the fundamentals cross the trigger value y^*.

The test is conclusive only if the null is rejected, in which case multiple equilibria provide the only possible explanation for the attack in question. Failure to reject the null supports equilibrium uniqueness, but does not rule out equilibrium multiplicity either.

This claim is substantiated rigorously by the analysis of Section 9 but the intuition is straightforward: multiple equilibria are consistent with *any* time path of devaluation expectations. This can be seen in the right panel of Figure 2: a sunspot could lead to an abrupt jump in $g(y_t)$ when y_t is increasing, steady or even decreasing within the multiplicity region.

3 The Model

The model assumes a target zone regime and specifies the behaviour of interest and exchange rates in terms of two variables that fully describe the fundamental state of the economy. One of the state variables, y, was introduced in Section 2 and underpins the central authority's decision to preserve or modify the exchange rate regime: a decision which typically has long-lasting consequences. y comprises the domestic trade balance, unemployment and real exchange rates. The other fundamental is denoted by f and consists of variables that influence the day-to-day movements of the exchange rate in a particular target zone. These variables include the domestic output and money supply, and the foreign money supply, interest rate and price level. The shocks to y and f are likely to be imperfectly correlated, given that they affect differently the low- and high-frequency components of the exchange rate. More concretely, the exogenous state is assumed to be driven by the following mean-reverting processes

$$\begin{aligned} \Delta y_t &= \rho_y (\overline{y} - y_t) \Delta + \sigma_y W^y_{t+\Delta} \\ \Delta f_t &= \rho_f (\overline{f} - f_t) \Delta + \sigma_f W^f_{t+\Delta} \end{aligned} \tag{2}$$

where $\rho_y > 0$, $\rho_f > 0$, $\overline{y}$, $\overline{f}$, $\sigma_y > 0$ and $\sigma_f > 0$ are constants and

$$\begin{pmatrix} W^y_{t+\Delta} \\ W^f_{t+\Delta} \end{pmatrix} \sim i.i.d. \left(\begin{bmatrix} 0 \\ 0 \end{bmatrix}, \begin{bmatrix} \sqrt{\Delta} & \sigma_{yf} \\ \sigma_{fy} & \sqrt{\Delta} \end{bmatrix} \right)$$

The specification of the empirical test allows for working with a simplified version of the function $g(\cdot)$, which translates y into devaluation expectations. The null of the test is cast in terms of market behaviour *prior* to an attack, which corresponds to values of y_t *smaller* than y^* in Figure 2. In addition, the first-order effects of the nonlinearities in $g(\cdot)$ set in at the time of the attack, i.e. for values of y_t *bigger* than y^*. Consequently, a linear approximation to $g(\cdot)$ is general enough for the purposes of the analysis and, in order to keep the number of free parameters to a minimum, I henceforth impose[8]

[8] The empirical exercise below identifies devaluation expectations up to an affine transformation. A

$$g(y_t) = y_t \text{ for } y_t \leq y^* \tag{3}$$

Next, I specify the determination of the log of the exchange rate, s_t. Denoting the date-t log central parity by c_t and the deviation of the log exchange rate from the log central parity by x_t, one obtains $s_t = c_t + x_t$. The value of c_t is set by the central authority, whereas devaluation expectations $\frac{E_t(\Delta c_t)}{\Delta} \equiv g(y_t)$. For the purposes of the analysis, c_t is set to zero without loss of generality because the central parity is constant during the periods of interest to the empirical exercise. Following the target zone literature, I assume that the position of the exchange rate within the band depends on the rate's expected change and on the fundamentals, f

$$x_t = \frac{E_t(\Delta s_t)}{\Delta} + f_t = y_t + \frac{E_t(\Delta x_t)}{\Delta} + f_t \tag{4}$$

The equation is derived from the forward-looking Cagan model, in which Δ alone denotes the period's length and Δ in front of a variable indicates the variable's one-period-*ahead* change.[9] Relevant only for periods preceding a speculative attack, (4) incorporates the implication of (3): $y_t = \frac{E_t(\Delta c_t)}{\Delta}$. Throughout the analysis, I refer to $\frac{E_t(\Delta x_t)}{\Delta}$ as *depreci*ation expectations.

The model is closed by assuming that uncovered interest rate parity holds. Denoting the one-period interest differential by ι, this implies[10]

$$y_t + \frac{E_t(\Delta x_t)}{\Delta} = \iota_t \tag{5}$$

3.1 Discussion of the model

Expressions (2)-(5) provide a simplified model that is easily employed in an empirical analysis of the periods preceding a violent speculative attack on target zone regimes. The model makes a series of implicit assumptions, which I spell out and rationalise in this section.

Since the exchange rate is a forward-looking variable, its pre-attack level incorporates expectations of market developments at the time of and after the attack. Following

more general linear function in equation (3) would thus not enrich the analysis.

[9] Krugman (1991) is the first paper to make use of the Cagan model in order to study exchange rate dynamics under a target zone regime. The Cagan model imlies $x_t = \alpha\left(\frac{E_t(\Delta x_t)}{\Delta} + y_t\right) + f_t$, where α indicates the extent to which the exchange rate level depends on its own expected rate of change. I set $\alpha = 1$, which is a typical assumption of the target zone literature. The message of the empirical results is not influenced by allowing α to range from 0.5 to 2.

[10] The interest differential is equal to the domestic minus the foreign nominally riskless interest rate.

Bertola and Svensson (1993), I assume that (i) equations (3)-(5) characterise also the aftermath of an attack and (ii) a successful attack leads to a new target zone band with an unchanged width and does not affect the deviation of the exchange rate from the central parity. At the time of the attack, the assumption in (3) can no longer be maintained and equation (4) is thus to be regarded as a simplification of

$$x_t = g(y_t) + \frac{E_t(\Delta x_t)}{\Delta} + h(f_t; y_t) \tag{6}$$

Non-linearities in $h(\cdot;\cdot)$, synchronised with those in $g(\cdot)$ via the common argument y_t, would accommodate extraordinary interventions that stabilise the exchange rate in the face of mounting devaluation expectations.[11] Under the above assumptions regarding the post-attack periods and mild technical restrictions on $g(\cdot)$ and $h(\cdot;\cdot)$, equations (4) and (5) imply a valid first-order approximation to the *pre-attack* levels of the exchange and interest rates.[12]

Even though I do not model explicitly the boundaries of the target zone, I assume that they are never violated owing to adjustments of the fundamentals, as implied by expression (2). Part of these adjustments would be due to interventions of the central authority, which may affect the exchange rate either directly (via f) or by influencing devaluation expectations. Since f and y follow mean-reverting processes, values of $\overline{f}$ and $\overline{y}$ sufficiently close to zero imply that the exchange rate regime is sustainable in the long run.

The autoregressive (AR) specification in (2) suggests intramarginal interventions and is consistent with the robust empirical finding that, in target-zone regimes, the exchange rate tends to cluster towards the center of the band. The expression stands in contrast to models that allow for interventions only at the boundaries of the target-zone band and imply counterfactually that the exchange rate should be close to these boundaries most of the time. The assumption in (2) is further supported by Garber and Svensson (1995) who reach the conclusion that, when both intramarginal *and* at-the-boundaries interventions are allowed for, the latter type of intervention is likely to be of little empirical significance.

It should be noted that the specification in (2) rules out jumps in the fundamentals, which is in line with macroeconomic data observed prior to and during the speculative attacks that the paper examines. It is conceivable, however, that official foreign reserve

[11] When the exchange rate is being stabilised, an increase in devaluation expectations surfaces through the interest differential.

[12] The technical requirement is that the first two (partial) derivatives of $g(\cdot)$ and $h(\cdot;\cdot)$ exist on the entire range of the functions. This is consistent with equations (3) and (4).

flows, which are notoriously difficult to measure, support the regime for some time by insulating market prices from gradually mounting devaluation expectations. An abrupt drying up of the reserve flows would constitute a jump in the fundamentals that would unleash prices and thus make a gradual attack, expressed temporarily in unobserved quantities, appear as violent. The paper assumes away such a scenario which is most likely irrelevant for the empirical exercise: Eichengreen, Rose and Wyplosz (1996) provide evidence of massive reserve flows supporting ERM currencies *during* the attacks in 1992. As argued above, the latter type of intervention is accommodated by the generalised specification of the exchange rate in (6).

4 Equilibrium conditions

Ruling out irrational bubbles, equations (2)-(5) imply that the observable exchange rate and interest differential are linear functions of the two fundamental variables

$$x_t = \left(\frac{\rho_f \overline{f}}{1+\rho_f} + \frac{\rho_y \overline{y}}{1+\rho_y} \right) + \frac{1}{1+\rho_f} f_t + \frac{1}{1+\rho_y} y_t \tag{7}$$

$$\iota_t = \left(\frac{\rho_f \overline{f}}{1+\rho_f} + \frac{\rho_y \overline{y}}{1+\rho_y} \right) - \frac{\rho_f}{1+\rho_f} f_t + \frac{1}{1+\rho_y} y_t \tag{8}$$

Equations (7) and (8) are consistent with the observed regularity that, in target zones, the instantaneous correlation between the exchange and interest rates may be of either sign.

The equilibrium conditions are interpreted as follows. The positive coefficients of devaluation expectations in equations (7) and (8) reflect the fact that the exchange rate and the interest differential increase with the (current) expected rate of change of the exchange rate. Since the shocks to f are mean reverting, a higher f_t implies a lower depreciation rate $\frac{E_t(\Delta x_t)}{\Delta}$ and, due to uncovered interest parity, a negative coefficient of f_t in equation (8). Finally, f_t influences the exchange rate via two channels: (i) directly and (ii) via $\frac{E_t(\Delta x_t)}{\Delta}$. The two forces move x_t in opposite directions but the first one dominates: this leads to the positive coefficient of f_t in equation (7).

The mean-reversion parameters, ρ_f and ρ_y, are key for the empirical exercise because they relate the unobserved fundamentals to observable interest and exchange rates. The logic behind the role of these parameters in equation (7) is seen as follows. Faster reversion to the mean is tantamount to less persistence in the exogenous processes and leads to a smaller impact of current exogenous shocks on future exchange rates. In turn, this implies a smaller impact of current shocks on the *current* exchange rate which, as indicated by equation (4), is a forward-looking variable. As a result, $\frac{\partial x_t}{\partial f_t}$ and $\frac{\partial x_t}{\partial y_t}$ decrease respectively in ρ_f and ρ_y. The role of the mean-reversion parameters in equation (8) is rationalised similarly.

5 Data

The empirical exercise analyses separately the 1992 speculative attacks on the French Franc and the Italian Lira and thus uses one of the following sets of daily data at a time:[13]

- Spot/next bid interest rates on euro-currency deposits[14] denominated in Italian Liras and Deutsche Marks and mid-ecu Lira-Mark exchange rates. The time span of these data is October 10, 1990 to July 15, 1992.

- Spot/next bid interest rates on euro-currency deposits denominated in French Francs and Deutsche Marks and mid-ecu Franc-Mark exchange rates. The time span is January 1, 1990 to September 16, 1992.

The Deutsche Mark interest rates are subtracted from the French Franc and the Italian Lira interest rates in order to obtain the corresponding differentials. The empirical analysis makes use of the interest rate data only through these differentials.

With the use of daily short-term interest rates, I attain the best possible approximation to a setting in which prices remain fixed between the issue and maturity dates of financial instruments.[15] Such a simplified setting is a key implicit assumption of the model of Tarashev (2003), which underlies the empirical treatment of devaluation expectations in this paper. Allowing for first-order price changes during the life of the instruments results in the model generating counterfactual chaotic dynamics in foreign exchange markets.[16]

[13] The data are extracted from the electronic database of the BIS. The exchange rate data are recorded at 2:15 PM GMT+1, whereas the interest rate data are bid rates recorded at 11:45 AM GMT+1. Market conventions postulate that a foreign-exchange transaction is executed two business days after the date on which a given exchange rate is set. However, spot/next interest rates are determined and recorded two business days before they come in force. Thus, the data are synchronized and in line with the assumptions of the model. Thanks to Bill English for bringing up the issue and to Gabriele Galati and Gaston Wieder for clarifying it.

[14] Interest rates on euro-currency deposits are viewed as free of political risk.

[15] The approximation is better when the importance of intra-day trades is smaller.

[16] Financial prices change during the life of the corresponding instruments if the latter are bought and sold after issuance and prior to expiration. In such a case, traders' strategy depends on the expectations of *their own* future decisions. As demonstrated by Jeanne and Masson (2000), the resulting market behaviour engenders chaotic dynamics in a large class of models to which the framework of Tarashev (2003) belongs.

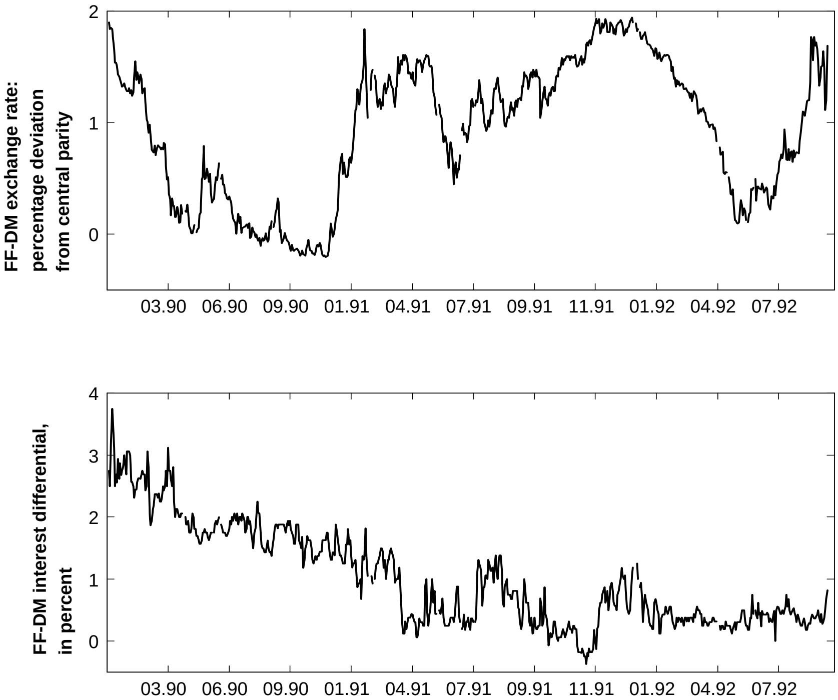

Figure 3: Data used for analysing the speculative attack on the French Franc

Short-term interest rates would in general be tightly linked to policy rates and thus might be expected to reflect market forces poorly. The concern does not seem to be borne out, however, because the employed series track closely the corresponding three-month interest rates, which are traditionally considered as reliable expressions of traders' beliefs.[17]

The interest rate data are filtered according to the criteria suggested by Rose and Svensson (1994). Particular attention is paid to spikes, which may be due to episodes of concerted unwinding of banks' foreign currency positions.[18] The filtered French Franc and Italian Lira data consist respectively of 684 and 444 exchange rates and interest rate differentials.

The last observation in a data set marks the end of a *pre*-attack period and is chosen as follows. In the spirit of Eichengreen, Rose and Wyplosz (1996), I identify

[17] From 1/1/1990 to 9/15/1992, the correlation between the French-Franc (respectively, Italian-Lira and Deutsche-Mark) spot/next and three-month interest rates is 0.82 (respectively, 0.85 and 0.93). Detrending the series via an HP filter with a smoothing parameter set to 14400, reduces the correlation coefficients to 0.55, 0.63 and 0.37.

[18] The filtering eliminated (only) the following two dates in *both* the Italian-Lira and French-Franc data sets: 12/27/90, 12/27/91.

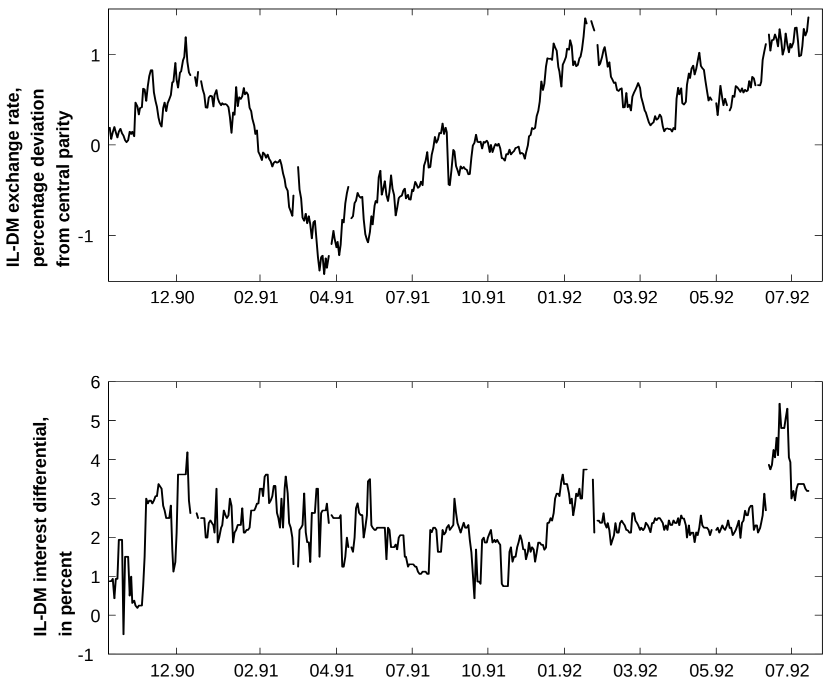

Figure 4: Data used for analysing the speculative attack on the Italian Lira

the violent speculative attack on each one of the two currencies with the first spike of the corresponding interest differential in 1992. I stop each sample two weeks before the relevant spike reaches its highest value: this incorporates the observed time between the inception and the peak of an attack.

The initial date of each sample is set as early as possible subject to excluding time periods during which non-modelled factors influence foreign exchange trading. The constraint leaves out data from the 1980s which witness a liberalisation of international financial markets and a substantial evolution of the credibility of ERM target-zone regimes. In addition, the Lira-Mark interest differential exhibits abnormal behaviour in August and September 1990, which prompts moving the start date of the Italian Lira data set until after that period.

Figures 3-5 allow for a visual assessment of the data. Figures 3 and 4 contain plots of the exchange rates and annualised interest differentials used in the empirical exercise. Figure 5 puts the interest differential in a time perspective. The series delimited by the vertical lines in that figure are exactly the same as the series in the bottom panels of Figures 3 and 4. I argue in the next section that the time limits on the data influence significantly the empirical procedure.

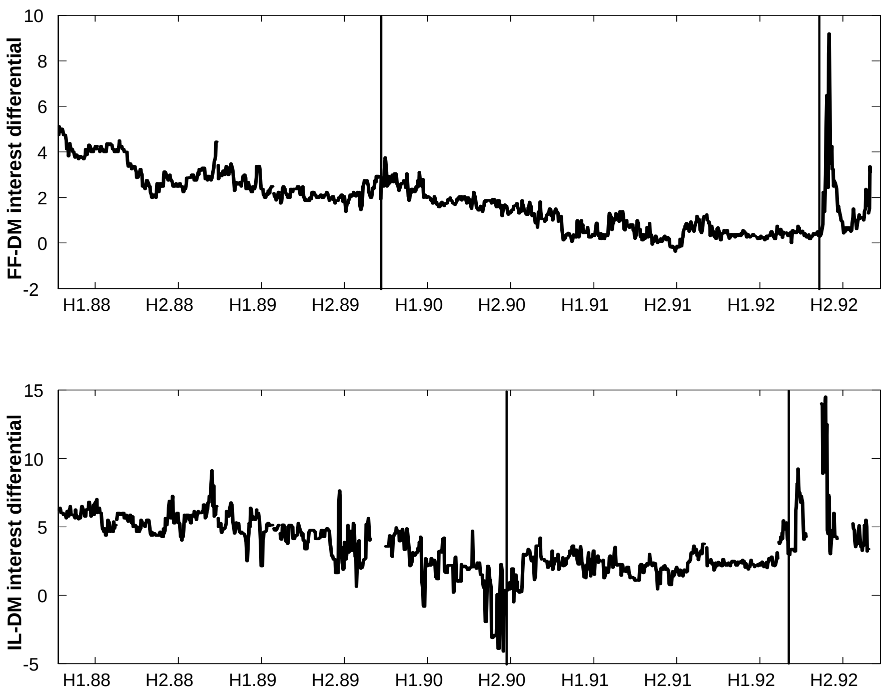

Figure 5: The interest differentials in a time perspective

6 Empirical Issues

I reasoned in previous sections that, in order to discover the type of equilibrium underlying devaluation expectations during a speculative attack, it is necessary to make inference about the mean-reversion parameters of the exogenous random variables, f and y. It turns out that the task cannot be fulfilled via standard inference techniques.

The mean-reversion parameters, ρ_f and ρ_y, capture *low*-frequency components of the exogenous variables and are thus poorly reflected in data with a limited time span. Crucially, sampling at a higher frequency does not solve the problem. The additional observations that such a sampling would bring in carry even less information about the mean-reversion parameters, especially when f and y are very persistent.

The above intuition is formalised by the analysis of Cavanagh (1985). For an illustrative example of some of the conclusions of that analysis, observe that assuming $\rho_y = \rho_f = \rho$ and adopting the definition $z_t \equiv \frac{1}{\sqrt{\Delta}} x_t$, expressions (2), (7) and (8) imply:

$$z_{t+\Delta} = \rho\left(\overline{f} + \overline{y}\right) + (1 - \rho\Delta)\, z_t + \xi_{t+\Delta} \qquad (9)$$

where the random variable $\xi_{t+\Delta} \sim i.i.d.\,(0,1)$. Using the term coined by Cavanagh

(1985), z follows a local-to-unity AR(1) processes. Recalling that Δ denotes the period's length, the sample size T is proportional to $\frac{1}{\Delta}$.[19] Thus, as the sample size increases, the AR coefficient $(1-\rho\Delta)$ and its estimator converge respectively to unity and the true value of the coefficient at the *same* rate, T. Consequently, even though $(1-\rho\Delta)$ is consistently estimated, the parameter ρ is *not consistently estimable.* In addition, the OLS estimator of ρ is biased upward and its asymptotic distribution is non-normal.

7 The Empirical procedure

This section lists the steps of a benchmark empirical procedure which (i) assumes that expression (2) reflects accurately the stochastic processes underlying the data and (ii) makes inference about the time profile of devaluation expectations. At the end of the section, I discuss a generalisation of the procedure, which accommodates richer features of the data.

The benchmark procedure consists of three steps. In the *first step*, I use equations (4) and (5) and data on exchange and interest rates in order to calculate the value of one of the exogenous fundamentals at each date in the sample: $f_t = x_t - \iota_t$. With the series $\{f_t\}_{t=1}^T$ at hand,[20] the *second step* relies on the bottom line of (2) in order to obtain the Dickey-Fuller statistic testing the null hypothesis $\rho_f = 0$.[21] For a particular value of the test statistic, Stock (1991) provides a 95% confidence interval (CI) for the value of ρ_f. Denoting a positive value on the CI by $\hat{\rho}_f$, in *step three* I calculate

$$\hat{y}_t = x_t - \frac{1}{1+\hat{\rho}_f} f_t = \frac{\hat{\rho}_f}{1+\hat{\rho}_f} x_t + \frac{1}{1+\hat{\rho}_f} \iota_t \tag{10}$$

In accordance with the theoretical model, negative or zero values of $\hat{\rho}_f$ are ignored at this step because they imply explosive fundamentals. Repeating step three at each date of the sample and for alternative values on the CI of ρ_f, I obtain a family of alternative series $\{\hat{y}_t\}_{t=1}^T$.

The empirical exercise tests the null hypothesis stated in Section 2, i.e. whether devaluation expectations increase to their highest pre-attack level just before the attack. For the purposes of such an objective, it is not necessary to estimate the time series $\{y_t\}_{t=1}^T$ but rather its time profile. In this respect, the series $\{\hat{y}_t\}_{t=1}^T$ suffices because

[19] The coefficient of proportionality is equal to the number of years in the sample. The broader is the time span of the sample, the less tightly is T linked to Δ and the less relevant are the issues discussed in the section.

[20] Henceforth, a variable in braces denotes a time series. Recall that the sample size T changes with the attack episode.

[21] Refer to Hamilton (1994) for a discussion of the Dickey-Fuller tests.

it consists of values of an *increasing* linear function of devaluation expectations: this is implied by equations (7) and (10) and the fact that $\frac{1}{1+\rho_y} > 0$.

Equation (10) shows that $\hat{\rho}_f$ influences the relative degree to which the dynamics of the exchange rate and the interest differential are imputed onto the estimated path of devaluation expectations. When the process of f is almost integrated, the time profiles of devaluation expectations and the interest differential are very similar. Symmetrically, the correlation between devaluation expectations and the exchange rate increases with the degree of mean reversion in f.

Observe that the adopted empirical approach makes direct inference regarding one of the parameters that cannot be estimated consistently, ρ_f, and eliminates the need to tackle the other one, ρ_y. The estimation uncertainty is thus reduced by taking advantage of (i) the null, which imposes only qualitative restrictions on devaluation expectations, and (ii) the flexibility to express equilibrium implications of the underlying model via different combinations of the observable variables.

In reality, the processes of the fundamentals need not be AR(1). The first step of the above procedure produces the time series $\{f_t\}_{t=1}^{T}$, which can be used to determine the order of the AR process of f that is supported by the data. The Box and Jenkins "modelling philosophy" provides a systematic way for doing so.[22] A more general process of f requires a generalisation of the remaining steps of the procedure, which is to also accommodate a richer process of y.

Such generalisations turn out to be warranted by the data used for analysing the speculative attacks on the French Franc and the Italian Lira. The augmented procedure is described in detail in Appendix 2 while its output is reported in Section 8 below.

Denoting the largest roots in the processes of f and y by $\left(1-\rho_f\Delta\right)$ and $\left(1-\rho_y\Delta\right)$, Appendix 2 demonstrates that the augmented procedure, just like the benchmark one, requires inference about only one of the two non-consistently estimable parameters: ρ_f. The parameters associated with the remaining roots of the two processes constitute the *raison d'être* of the augmented procedure but turn out to be estimated with negligible error. Since the inference about these parameters is trivial, I alleviate the exposition by referring only to the benchmark procedure when interpreting the empirical results.

8 Results

In this section I report and interpret time profiles of French Franc and Italian Lira devaluation expectations that are supported by the data at the 95% confidence level.

[22] Hamilton (1994) provides a description of the Box and Jenkins modelling philosophy.

The time profiles are used for testing the null hypothesis of equilibrium uniqueness.

8.1 The French Franc Episode

Figure 6 portrays an affine transformation of market expectations of a French Franc devaluation.[23] As motivated in Section 7, these plots represent only positive values in the 95% CI of ρ_f, i.e. values that imply mean reversion in f. Table 1 reports the underlying CI, which includes zero, together with the associated ADF statistic and the order of the AR process of f.

As argued in Section 7 the estimate of ρ_f determines the degree to which the characteristics of the two data series are imputed onto the estimated devaluation expectations. The plots in the top panel of the figure correspond to a stronger mean reversion in f and thus reflect predominantly the dynamics of the Franc-Mark exchange rate. In contrast, the plots in the bottom panel correspond to almost integrated processes of f and reflect to a larger extent the dynamics of the interest differential.

The key message of Figure 6 is that devaluation expectations do *not* reach their highest pre-attack level just before the attack. Namely, the peaks in the first quarter of 1990 and the first and fourth quarters of 1991 are higher than the peak at the end of the sample. Importantly, this is true for all of the examined time paths of devaluation expectations. When interpreted through the prism of the underlying theoretical model, these results suggest that the speculative attack on the French Franc must be driven by sunspots. Section 9 below illustrates in a concrete example how multiple equilibria can account for the series in both panels of Figure 6.

****

TABLE 1

Characteristics of the process followed by f

	AR(n) $n = ...$	ADF statistic	95% CI for ρ_f	70% CI for ρ_f
"French Franc" data	5	1.6316	$(-3.93, +10.36)$	$(-1.92, +6.25)$
"Italian Lira" data	2	4.821	(22.87, ...)[24]	(31.27, ...)

[23] Representing affine transformations of y_t, the actual values of the series cannot be interpreted. The labelling of the vertical axis is thus omitted in order to avoid unnecessary confusion. The horizontal axis measures time. A similar comment applies to Figure 7 below.

[24] Stock (1991) does not report the upper bounds of the confidence interval implied by an ADF statistic equal to 4.821. Filling in the dots with any numbers bigger than the respective lower bound does not alter the paper's conclusions.

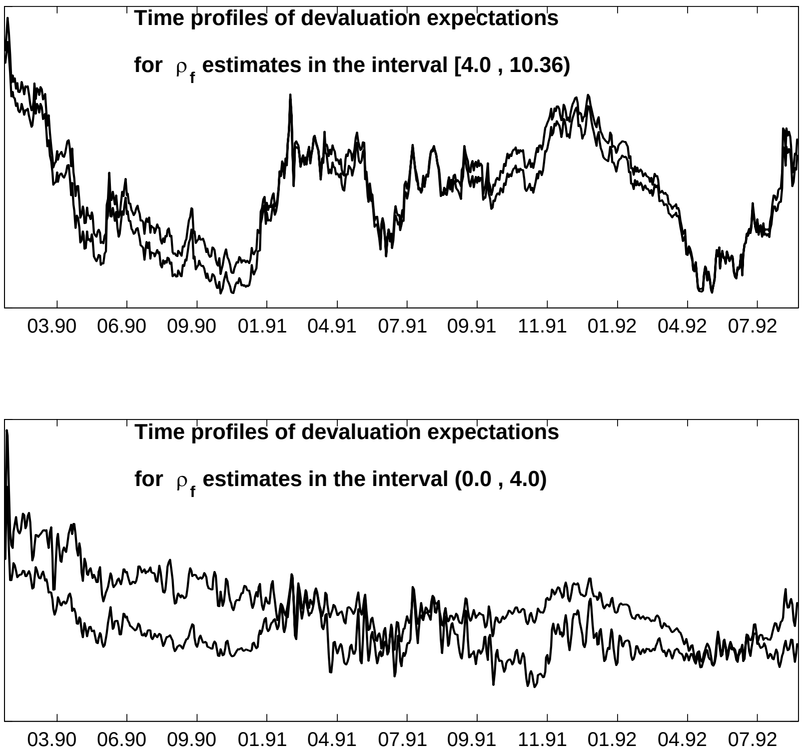

Figure 6: Run-up to the speculative attack on the French Franc

8.2 The Italian Lira Episode

Figure 7 portrays representative time profiles of Italian Lira devaluation expectations that are supported by the data at a confidence level of 95%. The plots are based on the 95% CI of ρ_f, which is reported in Table 1 and according to which the process of f is not integrated.

The figure indicates that devaluation expectations tend to increase over the last three quarters of the sample and settle at their highest pre-attack level in the last month of the sample. It is thus impossible to reject the hypothesis that the violent speculative attack on the Italian Lira is driven only by economic fundamentals in a unique equilibrium. The two peaks of devaluation expectations in December 1990 and February 1992 suggest that the attack is in the making for some time before it eventually erupts.

The small differences between the series in Figure 7 may seem surprising but can be rationalised by referring to equation (10), which implies that the uncertainty about ρ_f affects the uncertainty about devaluation expectations only via the coefficient $\frac{1}{1+\rho_f}$. In turn, the uncertainty about that coefficient decreases with the value at which the

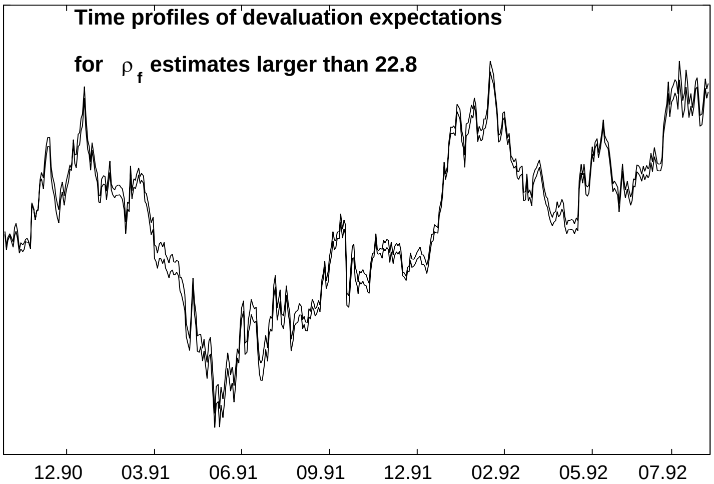

Figure 7: Run-up to the speculative attack on the Italan Lira

uncertainty about ρ_f is centered.[25] In the light of this fact, the high values in the CI of ρ_f (refer to Table 1) explain the similarity of the series in Figure 7.

9 A Target-zone model with equilibrium multiplicity

The analysis in the previous sections adopted the null hypothesis that speculative attacks erupt within equilibrium uniqueness. All along, I maintained that the alternative multiple-equilibria hypothesis cannot be ruled out regardless of whether the null is rejected or not. The claim is substantiated in this section which starts by modifying the framework in Section 3.

The building blocks of the model are as follows. As in Section 3, I denote by y the fundamentals affecting the central parity, c.[26] Next, I adopt a continuous-time setting and assume that, in the event of a parity realignment, there can only be a devaluation of a fixed size: $dc_t = 1$. Further, a devaluation on date t is possible only when two conditions are met: (i) $y_t > y^{cr}$ (a critical value of the fundamentals) and (ii) market devaluation expectations $\frac{E_t(dc_t)}{dt} \equiv g(y_t) > 0$. Under these conditions, a devaluation occurs over the next dt units of time with probability pdt.

The equilibrium devaluation expectations are then given by the following stylised version of the right panel in Figure 2:

[25] The claim is based on an application of the so-called Delta Method. See for example, Goldberger (1991), p. 102.

[26] Most of the notation in this section is borrowed from Section 3.

$$g(y_t) = \begin{cases} 0 & \text{if } y_t \le y^{cr} \\ 0 \text{ or } p & \text{if } y_t > y^{cr} \end{cases} \tag{11}$$

For $y_t > y^{cr}$, the equilibrium devaluation expectations are picked by a sunspot, which I assume to be independent of y_t. Let $\pi_1 dt$ denote the probability that, within the next dt periods, a sunspot shifts the economy from the "no-attack" state, in which $g(y_t) = 0$ to the "attack" state, in which $g(y_t) = p$. Likewise, if the peg is currently under attack, the instantaneous probability of shifting to the "no-attack" state is denoted by π_2. Even though the sunspot determines market devaluation expectations only when $y_t > y^{cr}$, it is observed for all y_t.

Finally, I specify the dynamics of the exchange rate, $s_t = c_t + x_t$. Since I am interested in pre-devaluation periods, I set $c_t = 0$ without loss of generality and postulate:

$$\begin{aligned} x_t &= g(y_t) + \frac{E_t(dx_t)}{dt} \\ dy_t &= \sigma_y dW_t^y \end{aligned} \tag{12}$$

where dW_t^y is a standard Wiener process. Expression (12) is a modified version of expressions (2) and (4). The exchange-rate fundamentals f are set to zero in this section because the main message can be conveyed by the impact of y_t on x_t.[27]

Expressions (11) and (12) fully specify the model, in which speculative attacks occur *only* under multiple equilibria. The equilibrium solution of x is derived in Appendix 3 to be a correspondence of y that is portrayed in Figure 8. The higher exchange rate schedule is associated with a sunspot state that would trigger an attack (a surge of devaluation expectations) when $y > y^{cr}$. The lower schedule corresponds to the sunspot state that keeps $g(y_t) = 0$ for all y_t.

At the boundary (y^{cr}) of the equilibrium-multiplicity region, there is a *discrete* change in the *expectation* of next period's *expected devaluation rate*. Nevertheless, no-arbitrage conditions require the path of the exchange rate to be continuous in each sunspot state. In the light of equation (12), this also requires the path of the expected depreciation rate to be continuous in each sunspot state. As demonstrated in Appendix 3, the no-arbitrage conditions result in two equilibrium values of the exchange rate for each value of the fundamental.

[27] As long as one is willing to consider only equilibrium solutions for x that are additively separable in f and y, setting the first variable to zero does not affect the analysis of the second variable's role. Of course, a realignment of the central parity is implemented by a discrete change in f; this is kept in the background of the current section.

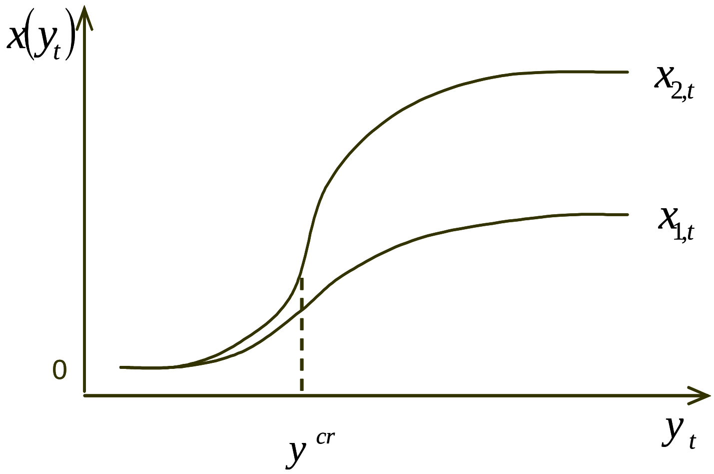

Figure 8: Equilibrium multiplicity under a target zone

For sufficiently strong fundamentals, the equilibrium exchange rate is virtually unaffected by the (remote) possibility of a surge in devaluation expectations. By (12), this translates into both exchange rate schedules converging to 0 when y_t decreases.

Symmetrically, for large values of y_t, the exchange rate is affected only by the sunspot and is higher in the "attack" sunspot state. Adopting the definitions $x_1^\infty \equiv \lim_{y_t \to \infty} x_{1,t}$ and $x_2^\infty \equiv \lim_{y_t \to \infty} x_{2,t}$, Appendix 3 shows that $\frac{x_{2,t}}{x_{1,t}} = \frac{1+\pi_1}{\pi_1} > 1$, which decreases in π_1, the probability of a switch to the "attack" state. If the exchange rate is to remain within a prespecified target zone band, the width of that band needs to be at least as large as x_2^∞.

We are now ready to see why a multiple equilibria hypothesis is irrefutable on the basis of devaluation expectations. Since the exchange-rate fundamentals f_t are set to zero, x_t can be interpreted in this section as the component of the exchange rate attributable to y_t, the fundamentals affecting the central parity. From such a point of view, Figure 8 portrays the multiple-equilibria counterpart of $\hat{y}_t$: the variable extracted from the data by filtering f_t out of the exchange rate on the premise of equilibrium uniqueness.[28] An increase of $\hat{y}_t$ to its highest pre-attack level just before the attack was interpreted in previous sections as evidence against the rejection of the equilibrium uniqueness hypothesis. Figure 8 illustrates, however, that the multiple-equilibria counterpart of $\hat{y}_t$ could exhibit a similar pattern: this would happen if y_t reaches its maximum value to the right of y^{cr} just before the first realisation of the "attack" sunspot state.

[28] Recall equation (10) for the definition of $\hat{y}_t$.

The figure also illustrates that no particular behaviour of devaluation expectations can be guaranteed in the presence of multiple equilibria. If $y_t > y^{cr}$, a sunspot may force $g(y_t)$ to jump from 0 to p even when y_t is decreasing. When y_t is decreasing, however, the component of the exchange rate, attributable to that variable, is also decreasing.[29]

10 Conclusion

The paper conducts an empirical analysis of two speculative attacks that occurred in the fall of 1992 and targeted respectively the French Franc and the Italian Lira. The empirical procedure makes inference about unobservable market devaluation expectations on the basis of interest and exchange rate data. This is accomplished via a theoretical model, which captures key stylized facts from target zone regimes and accounts for violent speculative attacks within both equilibrium uniqueness and multiplicity.

The empirical analysis relies explicitly on its theoretical foundations, which influence the choice of data. The adopted model of currency crises also highlights the fact that key estimators require a non-standard approach to statistical inference and underscores the extent to which identified empirical issues affect the conclusion of the analysis.

When interpreted through the prism of its theoretical foundations, the output of the empirical procedure suggests that the attack on the French Franc is driven by sunspots. In contrast, the data imply that the Italian Lira attack could be the result of either unique or multiple equilibria.

The empirical procedure is general enough to accommodate a wide range of speculative attacks on foreign exchange markets. Nevertheless, since there is abundant evidence that central authorities intervene actively during such episodes, data on official-reserve flows, if available at a sufficiently high frequency, would allow for a useful extension of the analysis.

[29] The analysis pays intentionally little attention to non-linearities in the relationship between fundamentals and observables prior to an attack. Reasons for the simplification were given in Section 3. In addition, at the paper's level of generality, pre-attack observable implications due to non-linearities under equilibrium uniqueness can always be matched under multiple equilibria via appropriate functional assumptions.

11 Appendix 1

In this appendix, I first outline key elements of the Tarashev (2003) model and then sketch its generalisation to a setting in which the exchange rate deviates from the central parity. Time subscripts are suppressed throughout the appendix, which uses notation from Sections 2 and 3.

The endogenous public signal (EPS) model of the above paper builds on the second-generation approach to currency crises. The model assumes a fixed exchange rate and allows market players to hold heterogeneous beliefs about the economic fundamentals. The optimal trading strategy of a speculator incorporates a subjective evaluation of the authority's incentives to abandon the peg *and* a private estimate of the beliefs of other traders. In equilibrium, market-wide devaluation expectations are aggregated in $g(y)$, where y reflects imperfectly the underlying fundamentals owing to *infinitesimal* informational noise. When $g(y)$ is used as an *input* to the analysis, as done in this paper, y can be treated as a fundamental variable without loss of generality.

The framework in Section 3 is based on an extension of the EPS model. A key implicit assumption of that framework is that the authority's decision to devalue is based on exogenous fundamentals and on market devaluation expectations but is independent of both the actual and expected dynamics of the exchange rate within the target zone band. The assumption is motivated by the following reasoning. When the regime is a target zone, deviations from the central parity are transitory while a change of the parity is typically irreversible in the short term. Consequently, *depreci*ation expectations play a smaller role than *devalu*ation expectations for the (non-modelled) long-term objectives of the private sector.[30] As a result, the former expectations have a smaller impact on the authority's objectives; for simplicity, the impact is assumed away.

The above assumption has three key implications, owing to which the new "target zone" features of the framework in Section 3 do not affect the determination of $g(y)$. First, the authority revises the central parity only on the basis of the factors underpinning its decision in the EPS model. Second, market devaluation expectations are exogenous to the dynamics of the exchange rate within a given target zone band. Third, the equilibrium beliefs about the likelihood of a devaluation constitute the only element of heterogeneity among speculators and are determined as in the EPS model. The last implication is due to the fact that the dynamics of the exchange rate within a given band are driven by f and y: variables that are publicly known by speculators who observe exchange and interest rates and know the parameters of the model.

[30] See Section 3 for a definition of the two types of expectations.

12 Appendix 2

The appendix is organised as follows. In Section 12.1, I introduce notation and state results that are used in a generalised empirical procedure deriving an affine transformation of the series $\{y_t\}_{t=1}^T$. The procedure is spelled out in Section 12.2. Finally, Section 12.3 provides specific results from applying the procedure to the French Franc and Italian Lira data sets described in Section 5.

12.1 Incorporating higher order AR processes of f and y

Let f and y follow respectively AR(p) and AR(n) processes whose largest roots are respectively $c_f \equiv \left(1 - \rho_f \Delta\right)$ and $c_y \equiv \left(1 - \rho_y \Delta\right)$.[31] A generalization of (7) is then of the form

$$x_t = \kappa + F(L) f_t + Y(L) y_t \tag{13}$$

where κ is a constant and the polynomials in the lag operator, $F(L)$ and $Y(L)$, are respectively of order $p-1$ and $n-1$. Further, the coefficients of $F(\cdot)$ and $Y(\cdot)$ are functions of the AR coefficients of f and y, respectively.

Defining y_t^f as:

$$y_t^f \equiv x_t - F(L) f_t \tag{14}$$

equation (13) implies that an affine transformation of y_t is given by

$$y_t^f \; Y^{-1}(L) \tag{15}$$

where y^f follows an ARMA$(n, n-1)$ process. The AR coefficients of the process of y^f are the AR coefficients of the process of y. The MA coefficients of the process of y^f correspond to the coefficients of $Y(L)$ up to a positive scalar factor.

Denote the AR and MA coefficients of the process followed by y^f respectively by b_q and μ_h, where $q \in \{0, 1, ..., n\}$ and $h \in \{0, 1, ..., n-1\}$. For $n \geq 2$, these coefficients are related in the following linear way

$$\mu_0 = 1, \text{ for } h \geq 1\text{: } \mu_h = \sum_{j=1}^{n-h} \left(\frac{1}{1+\Delta}\right)^j b_{j+h} \tag{16}$$

In the remainder of this section, I demonstrate that an estimate of an affine transformation of $\{y_t\}_{t=1}^T$ requires inference about exactly one non-consistently-estimable

[31] The assumption of a single local-to-unity root in the underlying processes of f and y is justified by the data.

parameter: ρ_f. In addition, the estimated transformation is seen as driven by an increasing function. As a by-product of the argument, I also show that y^f follows an ARMA$(n, n-1)$ process.

Equation (13) is underpinned by the generalised equilibrium condition:

$$x_t = \frac{\Delta}{1+\Delta} \sum_{s=t}^{\infty} \left(\frac{1}{1+\Delta}\right)^{\frac{s-t}{\Delta}} E_t \{f_s + y_s\} \tag{17}$$

where the summation is over t, $t+\Delta$, $t+2\Delta$, etc. If $p = 2$, (17) implies the following quasi-explicit version of (13):

$$x_t = \kappa + \frac{1}{[\Delta + (1-\beta_f)]\left(1+\rho_f\right)} [(\Delta+1) f_t - \beta_f c_f f_{t-\Delta}] + Y(L) y_t \tag{18}$$

where β_f denotes the second (smaller) root in the process of f. Estimates of the parameters of the process followed by f allow for an immediate calculation of the series $\left\{\hat{y}_t^f\right\}_{t=1}^T$, where "hats" denote henceforth estimates. Just as in the benchmark procedure of Section 7, *the calculation requires inference about the mean-reverting parameter* ρ_f. Equation (18) generalises trivially for $p \geq 3$.

The scenario, in which y follows and AR(1) process is discussed in the main text. When y is AR(n) and $n \geq 2$, $y_t B(L)$ is white noise and B is of order n. Likewise, by the definition in expression (14), $\frac{y_t^f B(L)}{Y(L)}$ is also white noise. Since $y_t Y(L)$ is the date-t forecast of future values of y, the coefficients in $Y(L)$ are functions of the coefficients in $B(L)$ and Y is of order $n-1$, implying that y^f follows an ARMA$(n$,$n-1)$.

The coefficients in $Y(L)$ are equal to the MA coefficients in the process of y^f (denoted in expression (16) by μ) up to a scalar factor, a division by which sets to unity the leading coefficient in $Y(L)$. That factor is always positive: by equation (4), a parallel upward shift of the historical time path of y increases x_t for all t. Thus, by (15), estimates of μ_h, for $h \in \{0, 1, ..., n-1\}$, are sufficient for an estimate of the values of an increasing linear function of the series $\{y_t\}_{t=1}^T$. Since the b parameters in expression (16) are consistently estimable, so are the μ parameters, which by deduction implies that it is *unnecessary to make inference about the second mean-reversion parameter* ρ_y.

12.2 An Outline of the generalised estimation procedure

Let f and y follow respectively AR(p) and AR(n) processes. This section describes the procedure for obtaining a point-wise estimate of devaluation expectations.

1. Derive $\{f_t\}_{t=1}^T$ on the basis of equations (4) and (5) and data on ι_t and x_t.

2. Using $\{f_t\}_{t=1}^T$, invoke the Box and Jenkins modelling philosophy to determine the order, p, of the process of f. Then, estimate the following Dickey-Fuller (DF) regression

$$f_t = \psi_0 + \psi_1 f_{t-\Delta} + \psi_2 \left(f_{t-\Delta} - f_{t-2\Delta}\right) + ... + \psi_p \left(f_{t-(p-1)\Delta} - f_{t-p\Delta}\right) + \eta_t \quad (19)$$

where η_t is a martingale difference sequence.[32] Construct the augmented Dickey-Fuller (ADF) statistic testing the hypothesis that $\rho_f = 0$. Using that statistic, refer to Stock (1991) for a confidence interval (CI) of ρ_f. To proceed, pick a point estimate of ρ_f from a grid on the CI.

3. Use the estimate of ρ_f from Step 2 and $\left(\hat{\psi}_2, \hat{\psi}_3, ..., \hat{\psi}_p\right)$ to derive $\hat{F}(L)$ and then obtain $\left\{\hat{y}_t^f\right\}_{t=1}^T$ on the basis of (14).

4. Invoke again the Box and Jenkins modelling philosophy in order to determine the order of the ARMA$(n, n-1)$ process followed by $\hat{y}^f$.[33]

   (a) If $n = 1$, $\left\{\hat{y}_t^f\right\}_{t=1}^T$ is an affine transformation of $\{y_t\}_{t=1}^T$ and the procedure stops.

   (b) If $n \geq 2$, go to Step 5.[34]

5. At this step, $\hat{y}^f$ follows an ARMA$(n, n-1)$ process with $n \geq 2$. Use a Wald test of the null hypothesis that the model-implied linear relationships, summarised in expression (16), are supported by the data.

   (a) If the null is not rejected, use n from Step 4 in order to deduce the MA parameters of $\hat{y}^f$. Derive an affine transformation of the y-series by estimating the AR parameters of $\hat{y}^f$ and then applying (15).

   (b) Different values of n are in order if (i) Step 5 is reached and (ii) the parameter relationships are *not* supported by the data for *all* values of ρ_f belonging to the CI in Step 2.[35]

[32] Note that this allows for GARCH elements in the error term.

[33] The data *in*variaby fail to reject the hypothesis that $\hat{y}^f$ follows an ARMA$(n,n-1)$.

[34] When $n \geq 2$, I adopt for the next steps an ARMA$(n, n-1)$ specification that is of the lowest order not rejected by the data. This helps avoid overparametrising the process of f

[35] The scenario of Step 5.b. never emerges in the empirical exercises of the paper.

12.3 Applying the generalised procedure to the data: some results

This section complements the discussion in Section 8 by providing technical detail underpinning the plots in Figures 6 and 7.

When applying Step 2 of Section 12.2 to the French Franc data set, I also determine the jointly normal asymptotic distribution of the estimators of $(\psi_2, \psi_3, ..., \psi_p)$. In large samples, it is safe to treat the CI of ρ_f as independent of the latter estimators.

The execution of Step 3. is based on a grid covering only *positive* values of the 95% CI of ρ_f (recall Table 1). For each point on that grid, I draw from the joint distribution of the estimators of $(\psi_2, \psi_3, ..., \psi_p)$ in order to construct a series $\left\{\hat{y}_t^f\right\}_{t=1}^T$. With each so-constructed series $\left\{\hat{y}_t^f\right\}_{t=1}^T$, I carry out Steps 4 and 5 to estimate the time profile of devaluation expectations.

For each value on the grid of ρ_f estimates, Steps 3 to 5 are repeated for 2000 draws from the joint distribution of the estimators of $(\psi_2, \psi_3, ..., \psi_p)$. As a by-product of this exercise, it turns out that the uncertainty about ρ_f is at the root of virtually all the uncertainty about y.

Figure 6 provides a representative picture of the time profiles of devaluation expectations supported by the data. The top panel of that figure is constructed with large estimates of ρ_f for which $\hat{y}^f$ follows an ARMA$(1,0)$ process. The bottom panel of the figure corresponds to smaller estimates of ρ_f for which $\hat{y}^f$ follows ARMA$(3,2)$. For the smaller estimates of ρ_f, I test the model as suggested in Step 5 of Section 12.2. The null hypothesis that the model is correct is never rejected: the p-values vary between 0.2 and 0.3.

Features of the Italian Lira data facilitate the execution of the steps listed in Section 12.2. In general, I proceed as with the French Franc data but make the following changes. For the execution of Step 3, I use a grid over the *entire* 95% CI of ρ_f because the interval includes only positive numbers. At different executions of Step 4, I find that the estimated series $\left\{\hat{y}_t^f\right\}_{t=1}^T$ supports invariably ARMA(1,0) processes, which prompts stopping the procedure at Step 4.a.

13 Appendix 3

This appendix is based on the model in Section 9 and derives the relationship between x_t and y illustrated in Figure 8. Recalling that the bottom schedule in the figure is denoted by $x_1(\cdot)$ and the top one by $x_2(\cdot)$, the equations behind these schedules are:

- for $y \leq y^{cr}$:

$$\begin{aligned} x_1(y) &= A_1 e^{\lambda_1 y} + \frac{\pi_1}{\pi_2} A_2 e^{\lambda_2 y} \\ x_2(y) &= A_1 e^{\lambda_1 y} - A_2 e^{\lambda_2 y} \end{aligned} \tag{20}$$

for $y \geq y^{cr}$:

$$\begin{aligned} x_1(y) &= p\frac{\pi_1}{1+(\pi_1+\pi_2)} + B_1 e^{-\lambda_1 y} + \frac{\pi_1}{\pi_2} B_2 e^{-\lambda_2 y} \\ x_2(y) &= p\frac{1+\pi_1}{1+(\pi_1+\pi_2)} + B_1 e^{-\lambda_1 y} - B_2 e^{-\lambda_2 y} \end{aligned} \tag{21}$$

where $\lambda_1 = \sqrt{\frac{2}{\sigma_y^2}}$, $\lambda_2 = \sqrt{\frac{2(1+(\pi_1+\pi_2))}{\sigma_y^2}}$ and the vector $\{A_1, A_2, B_1, B_2\}$ is determined by the four conditions:

1. $x_1(y)$ is continuous at $y = y^{cr}$
2. $x_2(y)$ is continuous at $y = y^{cr}$
3. $E_t\left(\frac{dx_1(y)}{dt}\right)$ is continuous at $y = y^{cr}$
4. $E_t\left(\frac{dx_2(y)}{dt}\right)$ is continuous at $y = y^{cr}$

Taking limits, we obtain the asymptotic values

$$\lim_{y \to -\infty} x(y) = 0 \tag{22}$$

$$\lim_{y \to \infty} x_1(y) = p\frac{\pi_1}{1+(\pi_1+\pi_2)} \tag{23}$$

$$\lim_{y \to \infty} x_2(y) = p\frac{1+\pi_1}{1+(\pi_1+\pi_2)} \tag{24}$$

To derive (20)-(21), I first record that[36]

$$x_1 = (x_2 - x_1)\pi_1 + E_t\left(\frac{dx_1}{dt} \middle| \begin{array}{l} \text{exchange rate} \\ \text{stays on } x_1 \text{ branch} \end{array}\right) \tag{25}$$

[36] In order to reduce the notational clutter, I keep the dependence of x_1 and x_2 on y in the background and suppress time subscripts when this creates no confusion.

whereas:

$$x_2 = \begin{cases} (x_1 - x_2)\,\pi_2 + E_t\left(\frac{dx_2}{dt}\middle| \begin{array}{l}\text{exchange rate}\\ \text{stays on } x_2 \text{ branch}\end{array}\right) & \text{for} \quad y \leq y^{cr} \\ (x_1 - x_2)\,\pi_1 + p + E_t\left(\frac{dx_2}{dt}\middle| \begin{array}{l}\text{exchange rate}\\ \text{stays on } x_2 \text{ branch}\end{array}\right) & \quad\quad\; y \geq y^{cr} \end{cases} \tag{26}$$

Below I solve for x_1 and x_2 when $y \geq y^{cr}$, which produces (21). Expression (20) is obtained similarly.

Adopting $\beta \equiv \frac{\sigma_y^2}{2}$ and assuming that $y \geq y^{cr}$, equations (25) and (26) can be rewritten as follows:

$$\begin{aligned} \begin{bmatrix} x_1'' \\ x_2'' \end{bmatrix} &= \frac{1}{\beta}\begin{bmatrix} 1+\pi_1 & -\pi_1 \\ -\pi_2 & 1+\pi_2 \end{bmatrix} * \begin{bmatrix} x_1 \\ x_2 \end{bmatrix} + \begin{bmatrix} 0 \\ -\frac{p}{\beta} \end{bmatrix} \\ X'' &= MX + V \end{aligned}$$

where x_1'', respectively x_2'', is the second derivative of x_1, respectively x_2, with respect to y.

Defining

$$\Xi \equiv \begin{bmatrix} 1 & \frac{\pi_1}{\pi_2} \\ 1 & -1 \end{bmatrix} = \left(\Xi^{-1}\right)^{-1} = \left(\frac{1}{1+\frac{\pi_1}{\pi_2}}\begin{bmatrix} 1 & \frac{\pi_1}{\pi_2} \\ 1 & -1 \end{bmatrix}\right)^{-1}$$

it follows that

$$D = \Xi^{-1}M\Xi = \begin{bmatrix} \frac{1}{\beta} & 0 \\ 0 & \frac{1+(\pi_1+\pi_2)}{\beta} \end{bmatrix}$$

where the non-zero entries of D are the eigenvalues of M and the columns of Ξ are the corresponding eigenvectors. An eigenvalue ψ solves $\det(M - \psi I) = 0$ and an eigenvector φ solves $(M - \psi I)\,\varphi = 0$.

Define next $\Theta \equiv \begin{bmatrix} \theta_1 \\ \theta_2 \end{bmatrix} = \Xi^{-1}X$ and observe:

$$\begin{aligned} \Xi^{-1}X'' &= \Xi^{-1}M\Xi\Xi^{-1}X + \Xi^{-1}V \\ \Theta'' &= D\Theta + \Xi^{-1}V \end{aligned}$$

Thus,

$$\theta_1'' = \frac{1}{\beta}\theta_1 + \left(\Xi^{-1}V\right)_{1,1}$$
$$\theta_2'' = \frac{1+(\pi_1+\pi_2)}{\beta}\theta_2 + \left(\Xi^{-1}V\right)_{2,1}$$

These are second-order differential equations that can be solved one at a time. Having determined Θ, premultiply by Ξ to find X, which produces expression (21). Expression (21) is obtained similarly and it remains to pin down the coefficients A_1, A_2, B_1 and B_2.

Suppose that x is continuous at y^{cr} and consider $E_t\left(\frac{dx_1}{dt}\right)$ for $y < y^{cr}$, where $E_t\left(\frac{dx_1}{dt}\right)$ stands for the expected rate of change of x given that the exchange rate is currently on the x_1 branch:

$$\begin{aligned}
E_t\left(\frac{dx_1}{dt}\right) &= (x_2 - x_1)\,\pi_1 + E_t\left(\frac{dx_1}{dt}\Big| \begin{array}{l}\text{exchange rate}\\ \text{stays on } x_1 \text{ branch}\end{array}\right) \\
&= \pi_1\left(-1-\frac{\pi_1}{\pi_2}\right)A_2 e^{\lambda_2 y} + A_1 e^{\lambda_1 y} + \frac{\pi_1}{\pi_2}\left(1+(\pi_1+\pi_2)\right)A_2 e^{\lambda_2 y} \\
&= A_1 e^{\lambda_1 y} + \frac{\pi_1}{\pi_2} A_2 e^{\lambda_2 y} = x_1
\end{aligned}$$

Now, consider $E_t\left(\frac{dx_1}{dt}\right)$ for $y > y^{cr}$:

$$\begin{aligned}
E_t\left(\frac{dx_1}{dt}\right) &= (x_2 - x_1)\,\pi_1 + E_t\left(\frac{dx_1}{dt}\Big| \begin{array}{l}\text{exchange rate}\\ \text{stays on } x_1 \text{ branch}\end{array}\right) \\
&= \pi_1 \frac{p}{1+(\pi_1+\pi_2)} \\
&\quad + \pi_1\left(-1-\frac{\pi_1}{\pi_2}\right)B_2 e^{-\lambda_2 y} + B_1 e^{-\lambda_1 y} + \frac{\pi_1}{\pi_2}\left(1+(\pi_1+\pi_2)\right)B_2 e^{-\lambda_2 y} \\
&= p\frac{\pi_1}{1+(\pi_1+\pi_2)} + B_1 e^{-\lambda_1 y} + \frac{\pi_1}{\pi_2} B_2 e^{-\lambda_2 y} = x_1
\end{aligned}$$

Thus, as long as x_1 is continuous at y^{cr}, $\lim_{y\to y^{cr}-} E_t\left(\frac{dx_1}{dt}\right) = \lim_{y\to y^{cr}+} E_t\left(\frac{dx_1}{dt}\right)$.

Let the intersection of the two parts of the x_1 schedule occur at the point (y^{cr}, x_1^{cr}). Note that $x_1^{cr} \in \left[0, p\frac{\pi_1}{1+(\pi_1+\pi_2)}\right]$, where the endpoints of the interval correspond to the two asymptotic values of the x_1 schedule: recall (22) and (23).

Even though the time derivative of x_1 is not well-defined at y^{cr}, one could write the expected rate of change of x_1 at $y = y^{cr}$ as

$$\lim_{y \to y^{cr-}} E_t \left(\frac{dx_1}{dt} | dy < 0 \right) \Pr(dy < 0) + \lim_{y \to y^{cr+}} E_t \left(\frac{dx_1}{dt} | dy > 0 \right) \Pr(dy > 0)$$

Moreover, the relationships in expressions (27) and (28) below are true whenever $\lim_{y \to y^{cr-}} E_t \left(\frac{dx_1}{dt} \right) = \lim_{y \to y^{cr+}} E_t \left(\frac{dx_1}{dt} \right)$.

First, we establish that it is not possible to have $x_1^{cr} = 0$:

$$\begin{aligned} \lim_{y \to y^{cr-}} E_t \left(\frac{dx_1}{dt} | dy < 0 \right) &= 0 = \lim_{y \to y^{cr-}} E_t \left(\frac{dx_1}{dt} \right) \\ \lim_{y \to y^{cr+}} E_t \left(\frac{dx_1}{dt} | dy > 0 \right) &> \lim_{y \to y^{cr+}} E_t \left(\frac{dx_1}{dt} \right) = 0 \end{aligned} \tag{27}$$

Likewise, $x_1^{cr} = p \frac{\pi_1}{1 + (\pi_1 + \pi_2)}$ is not possible because:

$$\begin{aligned} \lim_{y \to y^{cr-}} E_t \left(\frac{dx_1}{dt} | dy < 0 \right) &< \lim_{y \to y^{cr-}} E_t \left(\frac{dx_1}{dt} \right) = 0 \\ \lim_{y \to y^{cr+}} E_t \left(\frac{dx_1}{dt} | dy > 0 \right) &= 0 = \lim_{y \to y^{cr+}} E_t \left(\frac{dx_1}{dt} \right) \end{aligned} \tag{28}$$

Refer to equations (20)-(21) and note that the x_1 schedule to the left of (y^{cr}, x_1^{cr}) is convex whereas the right part is concave. Thus,

$$\begin{aligned} \frac{\partial}{\partial x_1^{cr}} \left\{ \lim_{y \to y^{cr-}} \left[E_t \left(\frac{dx_1}{dt} | dy < 0 \right) - E_t \left(\frac{dx_1}{dt} \right) \right] \right\} &< 0 \\ \frac{\partial}{\partial x_1^{cr}} \left\{ \lim_{y \to y^{cr+}} \left[E_t \left(\frac{dx_1}{dt} | dy > 0 \right) - E_t \left(\frac{dx_1}{dt} \right) \right] \right\} &< 0 \end{aligned} \tag{29}$$

By expressions (27)-(29), the following expression

$$\begin{aligned} & \lim_{y \to y^{cr-}} E_t \left(\frac{dx_1}{dt} | dy < 0 \right) \Pr(dy < 0) + \\ + \quad & \lim_{y \to y^{cr+}} E_t \left(\frac{dx_1}{dt} | dy > 0 \right) \Pr(dy > 0) - \lim_{y \to y^{cr\pm}} E_t \left(\frac{dx_1}{dt} \right) \end{aligned} \tag{30}$$

decreases with x_1 and crosses zero (exactly once) at $x_1^{cr} \in \left(0, p \frac{\pi_1}{1 + (\pi_1 + \pi_2)} \right)$.

The continuity of $E_t \left(\frac{dx_1}{dt} \right)$ at $y = y^{cr}$ is attained if and only if the x_1 schedule passes through (y^{cr}, x_1^{cr}). An analogous argument derives a necessary and sufficient condition for the continuity of $E_t\left(\frac{dx_2}{dt}\right)$. These two conditions, together with the continuity of the x_1 and x_2 schedules, pin down the values of A_1, A_2, B_1 and B_2.

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