CURRENCY CRISSES AND UNCERTAINTY ABOUT FUNDAMENTALS

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Abstract

This paper uses data on the distribution of macroeconomic forecasts during the Asian crisis to show that uncertainty about fundamentals contributes to exchange rate pressures. We find that a greater dispersion of the forecasts increases exchange rate pressures when the mean forecast of fundamentals is “good,” but reduces them when the mean forecast is “bad.” These results are consistent with the theoretical framework of Morris and Shin (1998), which predicts that, as the precision of information diminishes, exchange rate pressures increase or fall depending on whether the public signal and the mean of the private signals are “good” or “bad.”

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1 Introduction

Whether uncertainty about fundamentals plays a role in currency crises is an issue with important implications for both the theoretical and the empirical literature in international finance. The matter is also critical for policy purposes. For example, if uncertainty about fundamentals increases the probability of a speculative attack, then exchange rate regimes will be more vulnerable in periods of greater uncertainty and policymakers should adjust their policies accordingly. Moreover, to the extent that public authorities control the precision of information about fundamentals, a relevant role of uncertainty may carry implications for the optimal degree of transparency, the disclosure policy, as well as the timeliness of data releases.

In this paper, we study the effect of uncertainty about fundamentals with a dataset that includes forecasts of key macro variables for six Asian countries gathered by Consensus Economics. Figures 1 and 2 show that during the Asian crisis not only expected GDP growth deteriorated, but also the growth outlook became more uncertain, with a large increase in the dispersion of forecasts. The question we address is whether the increase in uncertainty (Figure 2) played a role in determining exchange rate pressures that is additional to the deterioration of the mean of expected fundamentals (Figure 1). To answer this question, we proceed in three stages.

First, we present a theoretical model à la Morris and Shin (1998) that predicts a role for uncertainty about fundamentals in currency crises (Section 2.) In this model, private information about fundamentals accounts for the presence of distinct individual forecasts, while public information accounts for economic risk and imprecise official information policies. Our main theoretical contribution is to derive testable implications in terms of shares of speculators attacking the currency – a feature with a direct empirical counterpart in some measures of exchange rate pressures. The key testable implication of the model is that the share of attacking speculators changes with the precisions of public and private information in a way that depends on whether the public signal and the mean of the private signals are “good” or “bad.” Specifically, the model predicts that, when the public signal and the average private signal are both “good,” a reduction in their precisions increases the share of speculators attacking the currency; whereas a reduction in the precisions of “bad” public and private signals has an opposite effect.

1 The evolution over time of the mean and variance of other macro forecasts in the Consensus Economics dataset is similar to that of GDP.

2 Differences in individual forecasts can also emerge with identical information sets because of the strategic behavior of forecasters (Laster et al. (1999)). In this paper, we simply assume that different forecasts are a symptom of different information sets. We can show, however, that our key testable hypotheses would remain the same even if forecasters behaved strategically (see Section 3.)
The intuition is straightforward. The less precise is a “good” public signal, the less it will be taken into consideration in forming expectations, and the larger will be the number of attacking speculators. Similarly, the more disperse are the idiosyncratic private signals around a “good” mean (which, in the model, corresponds to the realized fundamental), the larger will be the number of speculators receiving signals “bad” enough to make them attack the currency.

Second, we rigorously relate the empirical variance of the forecasts to the precisions of public and private information by making consistent assumptions on the way forecasters make their predictions (Section 3.) While data limitations do not allow us to determine whether it is a change in the precision of public or private information that causes a change in the dispersion of the forecasts, we can show that: \( i \) the mean of individual forecasts improves when the public and average private signals improve; and \( ii \) the dispersion of individual forecasts increases when the precisions of public and private information diminish. These links between the Consensus data and the model’s parameters allow us to test whether the precision of information affects exchange rate pressures with a sign that reflects how “good” public and private signals are.

Third, we verify the model’s testable implications using data for six Asian countries (Thailand, South Korea, Indonesia, Malaysia, Singapore, and Hong Kong) over the period January 1995 - May 2001, for which Consensus Economics forecast data are available (Section 4.) As it is common in the literature, we exploit time-series information to test an essentially static game, implicitly assuming that the data come from a myopic, repeated play of the one-shot game. Our estimates show that the dispersion of GDP growth forecasts: \( i \) is the most significant variable in our regressions; and \( ii \) is most significant and provides the largest contribution to the overall goodness-of-fit measures when we interact it with the mean of individual forecasts. We also find that, when the mean GDP forecast is “good,” the dispersion of the forecasts increases exchange rate pressures, and vice versa. Extensive robustness checks confirm these findings.

The empirical analysis of this paper has several novel aspects. First, almost no other paper in the empirical currency crisis literature has made uncertainty about fundamentals its central focus.\(^3\) Second, the tight link between the theoretical and the empirical model and the utilization of forward-looking survey forecasts of fundamentals, rather than only current levels of fundamentals, distinguish our analysis from previous empirical studies.

\(^3\) The mainstream empirical literature on currency crises, including Eichengreen et al. (1996) and Kaminsky and Reinhart (1999), has generally neglected the role of uncertainty about fundamentals. The two main exceptions are: Hodrik (1989), who unsuccessfully used estimated conditional variances of money supply, industrial production, and consumer prices, to account for the dynamics of the forward exchange rate premium; and Kaminsky and Peruga (1990), who estimated a GARCH-in-Mean restricted VAR model.
Third, this paper can be considered the first empirical test of some key predictions of models of currency crises à la Morris and Shin.

2 Theoretical background

This section analyzes the role of uncertainty about fundamentals in a model à la Morris and Shin with private information and a unique equilibrium, and discusses how it differs from that in the rest of the literature on currency crises. While modifying some of the other existing models might yield similar predictions, the class of models à la Morris and Shin has several advantages, including that of matching well the data used in our empirical analysis. Specifically, the assumption of private information accounts for the dispersion of individual forecasts and generates empirically plausible equilibria in which only a fraction of speculators attacks the currency, with or without success. \(^4\) In addition, the presence of a unique equilibrium allows to perform rigorous comparative statics exercises that are not possible in multiple-equilibrium models à la Obstfeld (1994 and 1996).

In stochastic “first-generation” models of currency crises, the variance of fundamentals affects the probability of a speculative attack at each point in time (Flood and Garber (1984)). In most of these models, greater uncertainty about fundamentals tends to increase the probability of a speculative attack. In Goldberg (1991), for instance, domestic credit growth follows a random walk process with errors distributed as a displaced exponential with zero mean; if the variance of the errors is below an upper bound, greater uncertainty increases the probability of an attack. Similarly, Flood and Marion (2000) extend Flood and Garber’s results to show that an increase in the expected post-attack variance of the exchange rate may lead the economy into an attack equilibrium even if the first moment of fundamentals is consistent with a no-attack equilibrium. \(^5\) Conducting a comparative statics exercise on the model of Grilli (1986), it is possible to show that a higher variance of fundamentals increases the probability of an attack as long as fundamentals are “good” but it may reduce it when fundamentals are sufficiently “bad.” This result is reminiscent of ours, but it is obtained with a model which considers only imprecise public information.

“Second-generation” models of currency crises have paid less attention to the role of uncertainty about fundamentals. These models are usually complete information models in which only the mean of the fundamentals

\(^4\) In models with complete – or incomplete but only public – information, only equilibria in mixed strategies would be characterized by a fraction of speculators attacking the currency.

\(^5\) In the related literature on stochastic target zones, Dumas and Svensson (1994) show that when the variance of the fundamentals is larger, the expected survival time of a target zone is shorter. Similarly, Bartolini and Prati (1999) find that the benefits of soft exchange rate bands decline as the variance of fundamentals increases.
matters (see, for example, Obstfeld (1996)). An exception is Sbracia and Zaghini (2001) who build a second-generation model of currency crises with incomplete public information about fundamentals and show that an increase in the variance of public information can make a unique equilibrium with a speculative attack prevail in a range of parameters in which, for lower levels of variance, there would be multiple equilibria. Annex 1 shows that the testable implications from an extension of this model would be essentially the same as in our model; such extension is relevant because it shows that our predictions on the effects of the mean and variance of public information would hold also in the presence of multiple equilibria.

Following Morris and Shin (1998), several papers have considered models with incomplete public and private information about fundamentals. These models would yield multiple equilibria with complete information, but a unique equilibrium emerges when the private signal about the state of fundamentals is sufficiently precise relative to the public signal. Nevertheless, “coordination failures” still characterize this unique equilibrium because the entire structure of beliefs (including the precision of public and private information), and not only the level of fundamentals, determines whether an attack or a no-attack equilibrium prevails. Thus, even though there is a unique equilibrium, exchange rate pegs can collapse for values of fundamentals that would have been consistent with the peg if only speculators’ expectations had been different.

Some recent papers have examined the effect of changes in the precision of public or private information on the likelihood of a currency crisis in models à la Morris and Shin. Using a uniform distribution for noisy private signals, Heinemann and Illing (2002) prove that an increase in the precision of private information reduces the likelihood of a currency crisis. Morris and Shin (2004) question, however, the robustness of this result and, in a somewhat different framework, find that greater precision of information does not always attenuate the coordination problem faced by the speculators. Finally, Metz (2002) shows that the effect of the precision of information on the decision rule of the government has a sign that varies with the nature of the signal (“good” or “bad”) and is opposite for public and private information.6

In this paper, we extend Metz’s result in order to obtain predictions about the effect of the precision of information not only on the decision rule of the government but also on the share of speculators attacking the currency, which is the correct theoretical counterpart of the indices of exchange rate pressure that we use in the econometric analysis. This extension is critical because it shows not only that sign of this effect varies with the

6In a related framework, Corsetti et al. (2003) show that the presence of a “large” speculator makes “small” speculators more aggressive in their attacking strategy and that the strength of this effect depends on the relative precision of private information of large and small investors.
nature of the signal ("good" or "bad") but that – differently from the effect on the decision rule of the government – tends to be the same for public and private information.

2.1 Complete information model

As a benchmark, we first consider a complete information model. We assume a continuum of speculators in an economy characterized by a state of fundamentals $\theta$ that can take values over the real line $\mathbb{R}$, with $\theta = +\infty$ corresponding to a situation of “sound fundamentals.” We assume that public authorities ("the government") are pegging the exchange rate and that speculators decide whether or not to attack it. If a speculator attacks and the government abandons the peg, the speculator obtains $D - t$, with $D > t > 0$; when the attack is not successful, the speculator loses the transaction cost $t$.\footnote{Here we take $D$ constant. Assuming that $D$ depends on the level of fundamentals $\theta$ (as in Morris and Shin (1998)) does not alter the results of the complete and the incomplete public information games. The model with both public and private information, instead, becomes too complicated to be solved analytically.} If speculators refrain from attacking, they get 0. The government’s utility from defending the currency is increasing in the fundamental $\theta$, and decreasing in the share of speculators that attack the currency, denoted by $l$. Specifically, we assume that the government gets $\theta - l$, when the peg is maintained and zero when it is abandoned.

We consider a very simple two-stage game with complete information. In the first stage, speculators observe $\theta$ and simultaneously decide whether to attack the currency. In the second stage, the government – who knows $\theta$ – observes the share of speculators attacking the currency and decides whether or not to maintain the peg.

This game can be solved backward by finding the government’s optimal strategy, which is simply the function:\footnote{We assume – without altering the analysis – that the government chooses to abandon the peg when it is indifferent.}

$$
\psi(\theta, l) = \begin{cases} 
    \text{abandon, if } \theta \leq l \\
    \text{defend, otherwise} 
\end{cases}.
$$

Given $\psi$, the solution of the reduced-form game of speculators provides the tripartition of the space of fundamentals that characterizes second generation models of currency crises. Specifically, since $l \in [0, 1]$, we find that if the fundamental $\theta$ lies in:\footnote{Hereafter we restrict our attention to pure strategies.}

- $(\infty, 0] \implies$ there is a unique equilibrium: all agents attack the currency and the government devalues;

- $[0, 1] \implies$ there is a unique equilibrium: the government succeeds in maintaining the peg.

- $[1, \infty) \implies$ there is a unique equilibrium: the government fails to maintain the peg.
• $(0, 1] \implies$ there are multiple equilibria: agents can either attack the currency (and force a devaluation) or refrain from attacking (and allow the peg to be maintained);

• $(1, +\infty) \implies$ there is a unique equilibrium: all agents refrain from attacking and the government maintains the peg.

Hence, outside the interval $(0, 1]$, maintaining the currency peg is solely a function of the fundamental $\mu$. By contrast, when $\mu$ falls in $(0, 1]$ the outcome depends on which self-fulfilling equilibrium speculators coordinate: if speculators expect the exchange rate peg to fail, they attack the currency and force the government to devalue; if they expect the peg to hold, they do not attack the currency and allow the government to maintain the peg.

2.2 Incomplete public and private information

In this section, we analyze the previous model assuming that there is incomplete public and private information about fundamentals. There are several reasons to assume that agents have private information. A noisy private signal may represent discrepancies in how public information is interpreted by different speculators. Kandel and Pearson (1995) and Kandel and Zilberfarb (1999) find empirical support for such heterogeneous processing of public information. Costs of information acquisition may produce heterogeneity in speculators’ information sets, as documented by Kaufmann et al. (2000). In foreign exchange markets, international banks may also gather valuable bank-specific information from monitoring the activity of their customers.

In this paper, the term public information is not used as a synonym for official information (i.e., information provided by the authorities of a country or by other national or international bodies) but as the antonym of private information. Public information consists of signals on the level of fundamentals that are common (publicly observable) to all agents, whereas private information differs from agent to agent. As a consequence, an increase in the variance of the distribution of public information does not necessarily reflect noisier official information but it could also be due to greater uncertainty – common to all agents – about the economic outlook or, more generally, to greater economic risk.\(^\text{10}\) An implication of this approach is that, unlike Morris and Shin (2002), we cannot perform welfare analysis on the provision of public information.

\(^{10}\)The sharp increase in the dispersion of GDP forecasts in the aftermath of currency crises documented in Figure 2 may also reflect an increase in “model uncertainty” (i.e., an increase of the uncertainty about the “true” model of Asian economies), as defined by Routledge and Zin (2001). In the theoretical framework of our paper, an increase in model uncertainty may translate into a higher variance of public or private information, depending on whether uncertainty increased in a similar or different way across agents.
Virtually any event that is publicly observable and affects economic fundamentals – including a currency crisis elsewhere or rumors of political troubles – could be classified under the public information label. The crisis in Thailand, for example, may have made the growth outlook of other Asian countries equally more uncertain for all agents. At the same time, uncertainty about the policies that each country would follow in the midst of the crisis may well have contributed to the overall uncertainty.

In the incomplete information model we assume that speculators do not know the fundamental \( \theta \) but only have expectations about it, given by the following prior normal probability distribution

\[
\Theta \sim \text{Norm}(y, 1/\alpha),
\]

with the “public signal” \( y \in \mathbb{R} \) and the “precision of public information” \( \alpha > 0 \). Since \( \Theta \) is common knowledge to all speculators, this probability distribution represents the public information available to them. Suppose also that each speculator \( i \) receives a private signal \( x_i \) drawn from the following normal distribution

\[
X_i \mid \theta \sim \text{Norm}(\theta, 1/\beta),
\]

with \( X_i \) and \( X_j \) independent given \( \theta \) for each \( i \neq j \), and “precision of private information” \( \beta > 0 \). Note that by setting either \( \alpha = +\infty \) or \( \beta = +\infty \) (or both) we get back to the complete information model.

When private information is sufficiently precise with respect to public information, this model entails a unique equilibrium. As was first shown by Carlsson and van Damme (1993), this result is driven by the lack of common knowledge induced by the presence of private information. Appendix A.2 illustrates why the lack of common knowledge leads to a unique equilibrium. A condition that ensures the existence of a unique equilibrium is:

\[
\beta > \frac{\alpha^2}{2\pi}. \tag{1}
\]

The intuition for this condition is straightforward. If private signals were not sufficiently informative with respect to the public signal, speculators would

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\(^{11}\) Using a somewhat different framework, Chan and Chiu (2002) show that if the complete information game does not include two regions characterized by a unique equilibrium, then private information – no matter how precise – would not result in a unique equilibrium. In other words, to obtain the unique-equilibrium result, it is crucial that there be a non-negative probability of \( \theta \) belonging to \((-\infty, 0)\) and to \((1, +\infty)\). Our model fulfills this condition because we assume normal distributions. Angeletos et al. (2003) show that, if the monetary authority modifies interest rates in response to the realization of fundamentals (thus, revealing its information on it), common knowledge would be restored together with multiple self-fulfilling equilibria. In this paper, we rule out such endogenous interest rate policy or, equivalently, assume that exchange rate fundamentals are not the only factors influencing government’s decisions regarding interest rates, thus preventing interest rate policy from fully restoring common knowledge about exchange rate fundamentals.
regard them as unreliable and continue to ground their decisions mostly on public information, restoring a high degree of common knowledge and, thereby, making multiple equilibria possible.

Proposition 1 (Morris and Shin, 2002a; Metz, 2002) If \( \bar{\alpha} > \frac{\bar{\alpha}^2}{2\bar{\gamma}} \), there exists a unique equilibrium, consisting of: (i) a unique value of the private signal \( x^* \) such that each speculator receiving a signal lower than \( x^* \) attacks the currency peg; and (ii) a unique level of the fundamentals \( \theta^* \in [0, 1] \) such that the government abandons the peg when fundamentals are lower than \( \theta^* \).

Appendix A.1 proves Proposition 1. As a consequence of the unique equilibrium result, for given values of the parameters \( y, \alpha, \) and \( \beta \) the survival of a currency peg depends solely on the actual fundamental \( \theta \). However, the equilibrium trigger points \( \theta^* \) and \( x^* \) are both functions of \( y, \alpha, \) and \( \beta \). Thus, speculators’ expectations still matter, as changes in the public signal and in the precision of public and private information change the range of values of the fundamentals for which the peg is maintained. Note also that the existence of a unique equilibrium does not eliminate all the “inefficiencies” of the model: when \( \theta^* \in (0, 1) \) we can still have currency crises (for \( 0 < \theta < \theta^* \)) that could have been avoided with complete information and speculators coordinating on the good equilibrium.\(^{12}\)

The presence of a unique equilibrium allows for rigorous comparative statics. Specifically, by assuming that condition (1) holds, we can study the effects of the parameters \( y, \alpha, \) and \( \beta \) on \( \theta^* \) and \( x^* \). Most importantly, we can calculate the effects of the parameters on the probability that speculator \( i \) will attack, \( \Pr(X_i \leq x^* | \theta) \). This probability represents the share of speculators attacking the currency and, therefore, has an empirical counterpart in indices of exchange rate pressure.

The share of speculators attacking the currency is equal to \( \Pr(X_i \leq x^* | \theta) = \Phi \left( \sqrt{\beta}(x^* - \theta) \right) \). This probability depends on the actual fundamental \( \theta \) and on the parameters \( y, \alpha, \) and \( \beta \). Given the results in Appendix A.2, by differentiating \( \Phi \left( \sqrt{\beta}(x^* - \theta) \right) \) one can obtain:

Proposition 2 Assume that \( \beta > \frac{\bar{\alpha}^2}{2\bar{\gamma}} \); then the probability \( \Pr(X_i \leq x^* | \theta) \) is: (i) decreasing in \( \theta \); (ii) decreasing in \( y \); (iii) decreasing (increasing) in \( \alpha \) if \( y > s_1 \) (\( y < s_1 \)), where \( s_1 \in \mathbb{R} \); (iv) decreasing (increasing) in \( \beta \) if \( \frac{(x^*-\theta)}{2\sqrt{\beta}} + \sqrt{\beta} \frac{dx^*}{d\beta} < 0 \) \( \frac{(x^*-\theta)}{2\sqrt{\beta}} + \sqrt{\beta} \frac{dx^*}{d\beta} > 0 \).\(^{13}\)

\(^{12}\)Morris and Shin (2002a) further show that the unique Bayesian Nash equilibrium of this game is also the unique strategy profile that survives iterated deletion of dominated strategies. In a related framework, Heinemann and Illing (2002) exploit this property to show that the introduction of sunspots (correlation devices unrelated to fundamentals) does not restore multiplicity of equilibria.

\(^{13}\)The expressions of \( s_1 \) and \( \frac{dx^*}{d\beta} \) are derived in Appendix A.2.
An improvement in the actual fundamental $\theta$ or in the public signal $y$ reduces the share of speculators attacking the currency (points (i)-(ii) in Proposition 2). The intuition of these effects is straightforward and it is illustrated in Figure 3, which shows how, for given values of the parameters $\alpha$, $\beta$, and $y$, the share of attackers is a negative function of the realized fundamental $\theta$. The 45-degree line shows the minimum share of attackers that would make the government abandon the peg. The point at which the share of attacking speculators crosses the 45-degree line corresponds then to $\theta^*$ on the horizontal axis. The dashed line shows how $\theta^*$ and the share of attacking speculators falls as $y$ increases.

Key to understand the intuition of the effects of $\alpha$ and $\beta$ (points (iii)-(iv) in Proposition 2) is the role of coordination in currency crisis games. In deciding whether to attack the currency, speculators need to consider not only their own expectations about fundamentals, but also what other speculators expect about fundamentals, what other speculators expect about others’ expectations about fundamentals, and so on. These expectations depend on the parameters $\alpha$ and $\beta$, which can assume values that either strengthen or weaken the beliefs of each individual on the other speculators’ decision to attack the currency. For example, if one speculator expects others to have similar beliefs, he will be more inclined to act on them.

These beliefs about the beliefs of others depend on the ratio between the precision of the two signals, $\frac{\alpha}{\beta}$, because this ratio determines the relative weight assigned to public and private information in the posterior beliefs and, in turn, the extent to which individuals can expect their beliefs to be shared. When speculator $i$ receives a message $x_i$ his “ex-post expected fundamental” – i.e. the Bayesian update of the “ex-ante expected fundamental” $y$ – is:

$$f_i^*(x_i) = E[\Theta | x_i] = \frac{\alpha y + \beta x_i}{\alpha + \beta}.$$ 

(2)

Suppose that $y$ is sufficiently high (i.e., conditions (13)-(15) in Appendix A.2 hold) so that speculators expect, ex-ante, “good” fundamentals. In this situation, if the precision ratio $\frac{\alpha}{\beta}$ is also high, speculators know that other speculators have formed their ex-post expectations attributing a large weight to the “good” public signal $y$ and will be less inclined to attack the currency. As a result, speculators will be less aggressive. By contrast, if $\frac{\alpha}{\beta}$ is low, speculators will be less inclined to rely on the “good” public signal $y$, because they know that the others are assigning a large weight to their random private signals. In other words, coordination on a “good” public signal is more difficult when the random component $x_i$ in each individual expectation carries a large weight. The same reasoning applies when $y$ is “bad”.\(^{14}\)

\(^{14}\)Heinemann and Illing (2002) obtain a different result on the effect of private inform
When \( y \) is sufficiently good (bad), an increase in \( \alpha \) improves then the coordination of speculators and reduces (increases) the share of attacking speculators by decreasing (increasing) the equilibrium trigger levels \( \theta^* \) and \( x^* \). This effect is illustrated for a “good” \( y \) in Figure 4.

The effect of changes in the precision of private information \( \beta \) on the share of speculators attacking the currency is more complex because it depends not only on the public signal \( y \) but also on the realized fundamental \( \theta \), around which the signals \( x_i \) are centered. First, a higher \( \beta \) reduces the ratio \( \frac{\theta}{\beta} \) and makes it more difficult for speculators to coordinate on the public signal \( y \). For a sufficiently good (bad) \( y \), a higher \( \beta \) tends to increase (reduce) the share of attacking speculators by increasing (reducing) the equilibrium trigger levels \( \theta^* \) and \( x^* \). This effect is illustrated for a good \( y \) in Figure 5 by the shift from the thick solid line to the dotted line.

A higher \( \beta \), however, has also the second very important effect of concentrating the private signals around the actual fundamental \( \theta \). This second effect can either offset or reinforce the first, depending on the realization of \( \theta \). If the realized \( \theta \) is sufficiently “good”, the number of signals below the threshold \( x^* \) may diminish enough to more than offset the effect of a higher \( x^* \) (due to an increase in \( \beta \) when \( y \) is “good”) on the share of attacking speculators; as a result, the share of attacking speculators decreases following the increase in \( \beta \). Figure 5 illustrates the effect of a higher \( \beta \) on the dispersion of the private signals as a shift of the curve representing the share of attacking speculators from the dotted line to the thin solid line. Figure 5 also shows that, when the realized fundamental \( \theta \) is, instead, “bad”, the second effect reinforces the first and the share of attacking speculators increases.

When we consider the polar case in which \( y \) is “bad”, we have symmetric results. In this case, a higher \( \beta \) makes it more difficult to coordinate on the “bad” public signal and attack the currency, but, if the realized fundamental \( \theta \) is also “bad”, the effect of the greater concentration of the private signals around \( \theta \) may dominate and the share of speculators attacking the currency increase, just as in the case of an increase in \( \alpha \) with \( y \) “bad.”

We can then conclude that the effect of the precision of private information tends to be similar to that of public information, provided that the condition: in their model an increase in the precision of private information, \( \beta \), always decreases \( \theta^* \), making speculative attacks less likely. However, Heinemann and Illing assume that \( \theta \) is uniformly distributed over the unit interval. In the terms of our model, this assumption would correspond to a fixed \( y \), set equal to 1/2. Hence, their result is consistent with our model – which, for a fixed \( y \), predicts that an increase in \( \beta \) always reduces \( \theta^* \), provided that condition (14) is not fulfilled. It should also be noted that when uncertainty is high the model of Heinemann and Illing tends to favor the attack strategy profile because speculators’ payoffs, given a successful attack, are assumed to depend negatively on \( \theta \). This means that if the attack is successful and \( \theta \) is low, speculators get a large payoff, whereas they lose only the transaction cost \( t \) if the attack is not successful. As a result, in that model an increase in uncertainty – making extreme values of \( \theta \) more likely – tends to drive speculators to the attack strategy.
average private signal (i.e., the actual fundamental) and the public signal (which is the ex-ante expected fundamental) are either both sufficiently good or both sufficiently bad; otherwise, the effect of private information precision may be opposite to that of public information precision.

3 Testable implications

Bringing the model of Section 2.2 to the data presents several challenges, mainly associated with the fact that distinct empirical measures of \( \alpha, \beta, \theta \), and \( y \) are not available. We can verify, however, some of the model’s predictions by using forecasts of macroeconomic variables collected by Consensus Economics. In this section, we precisely relate the mean and the variance of the forecasts to the parameters of the theoretical model. Specifically, we show that the mean forecast is positively related to both the public signal \( y \) and the mean of the private signals \( \mu \) and that the variance of the forecasts is inversely related to the precisions of public and private information. These results will allow us to test whether: i) improvements in \( y \) and \( \mu \) are associated with lower exchange rate pressures; ii) \( \alpha \) and \( \beta \) have an effect on exchange rate pressures that is additional to that of \( y \) and \( \theta \); and iii) the effect of \( \alpha \) and \( \beta \) depends on how “good” \( y \) and \( \theta \) are.

The Consensus Economics dataset gathers individual forecasts of economic variables (GDP, current account, inflation, ...) formulated by a set of professional forecasters. To relate these predictions to the theoretical model, we assume that each forecaster declares to Consensus Economics his ex-post expected fundamental. (If the forecasters had strategic objectives and chose their forecasts following any of the strategies considered in Ottaviani and Sorensen (2001), our testable implications would remain unchanged.\textsuperscript{15}) Recall that, given the message \( x_i \), the posterior probability distribution is

\[
\Theta \mid x_i \sim N \left( \frac{\alpha y + \beta x_i}{\alpha + \beta}, \frac{1}{\alpha + \beta} \right).
\]

Our assumption implies that the individual forecast that agent \( i \) (i.e., the agent receiving the message \( x_i \)) reports to Consensus Economics is the ex-post expected fundamental \( f^e_i (x_i) \), given by equation (2). Let us consider the mean of the individual forecasts, i.e.:

\[
f^e(x_1, ..., x_n) = \frac{\sum f^e_i (x_i)}{n} = \frac{\alpha}{\alpha + \beta} y + \frac{\beta}{\alpha + \beta} \sum x_i ,
\]

where \( n \) is the number of forecasters. Given the fundamental \( \theta \), for \( n \) that goes to \( +\infty \) this random variable converges to:

\[
f (\theta) = E [ f^e (X_1, ..., X_n) \mid \theta] = \frac{\alpha}{\alpha + \beta} y + \frac{\beta}{\alpha + \beta} \theta.
\]

\textsuperscript{15}A formal proof is available from the authors upon request.
Thus, if \( n \) is sufficiently large, by using the mean of the individual forecasts in the empirical analysis we use a variable that is influenced by \( \theta \) and \( y \). Recall that \( \theta \) and \( y \) have the same effects on the share of attackers: when \( \theta \) or \( y \) improve (deteriorate), pressures on the exchange rate will abate (strengthen). We can then use the mean forecast to test whether improvements in the public signal or in the realized fundamentals lead to a reduction in exchange rate pressures.

Note also that the theoretical model suggests that \( E[f(\Theta)] = y \); thus, on average the mean of individual forecasts is equal to the public signal \( y \) and does not depend in any systematic way on \( \alpha \) and \( \beta \). Similarly, in our empirical work we expect that, along the time-series dimension, the mean of individual forecasts does not depend on \( \alpha \) and \( \beta \).

The theoretical model also implies that the precisions of public and private information affect exchange rate pressures (points (iii)-(iv) in Proposition 2) and their effects are additional to that of \( \theta \) and \( y \) (points (i)-(ii) in Proposition 2). In other words, Proposition 2 suggests that even if \( \theta \) and \( y \) remain unchanged, speculative pressures on the exchange rate vary with the variance of public and private information. Empirically, changes in the precision of public and private information will be reflected in the variance of the individual forecasts:

\[
[\sigma^2(x_1, ..., x_n)]^2 = \sum \frac{[f^e_i(x_i) - f^e(x_1, ..., x_n)]^2}{n} = \frac{\beta^2}{(\alpha + \beta)^2} \frac{\sum (x_i - \overline{x})^2}{n},
\]

where \( \overline{x} = n^{-1} \sum x_i \). Given the fundamental \( \theta \), for \( n \) that goes to \(+\infty\) this random variable converges to:

\[
\sigma^2(\theta) = E[(\sigma^2(X_1, ..., X_n))^2 | \theta] = \frac{\beta}{(\alpha + \beta)^2}.
\]

Hence, for \( n \) sufficiently large, a change in \( y \) affects the mean of the individual forecasts \( f^e_i \) but does not affect their variance \( \sigma^2 \), which only depends on \( \alpha \) and \( \beta \). According to the model of Section 2.2, changes in the mean of the Consensus Economics forecasts shown in Figure 1 cannot explain coincident changes in the dispersion of the forecasts shown in Figure 2.

It is apparent from expression (5) that while an increase in \( \alpha \) always implies a decrease in \( \sigma^2 \), an increase in \( \beta \) does not necessarily reduce \( \sigma^2 \). This result is easily explained. On the one hand, \( \beta \) tends to reduce \( \sigma^2 \) as it decreases the dispersion of the messages \( x_i \). On the other hand, for given messages \( x_i \), the rise in \( \beta \) increases the weight of the private messages in the individual predictions (2), making them more heterogeneous between the forecasters. The first (second) effect dominates when \( \beta > \alpha \) (\( \beta < \alpha \)).

We conduct our empirical investigation on the assumption \( \beta > \max \left\{ \alpha, \frac{\alpha^2}{2\pi} \right\} \). The condition \( \beta > \alpha \) ensures that \( \sigma^2 \) is decreasing in \( \beta \), so that we can always interpret a decline in \( \sigma^2 \) as due to an increase in either \( \alpha \) or \( \beta \) or both.
The condition $\beta > \frac{\alpha^2}{2f}$ ensures the existence of a unique equilibrium and that Proposition 2 holds.\textsuperscript{16}

Proposition 2 implies that the effect of $\sigma^2$ on speculative pressures depends on how “good” $\theta$ and $y$ are. We therefore estimate a specification of the following general form:

$$ERP_t = \gamma_0 + \gamma_1 f^e_{t-1} + \gamma_2 \sigma^e_{t-1} \cdot (f^e_{t-1} - \gamma) + \gamma_3 e_{t-1} + \epsilon_t$$

where $ERP$ is a measure of exchange rate pressure, $f^e$ is the mean of the individual forecasts from (3), $\sigma^e$ is the standard deviation corresponding to the square root of (4), $\gamma$ is the estimated threshold separating “good” from “bad” means $f^e$, and $e$ is the real effective exchange rate. All regressors are lagged one period to avoid simultaneity bias.

We expect the coefficient $\gamma_1$ to be negative because an improvement in $\theta$ or $y$ eases the pressure on the exchange rate.

While there are good reasons to expect $\gamma_2$ to be positive (as we actually estimate it in Section 4), the possibility of an opposite sign cannot be ruled out. Proposition 2 implies that, when $y$ and $\theta$ are both “good”, a reduction in the precision $\alpha$ and/or in the precision $\beta$ (i.e., an increase in the dispersion of the individual forecasts, $\sigma^e$) would determine an increase in the share of speculators attacking the currency (i.e., an increase in exchange rate pressures, $ERP$). Given that $f^e$ increases with $y$ and $\theta$ because of (3), it is clear that a situation in which $y$ and $\theta$ are both “good” will be associated with a relatively high $f^e$ and an expected positive sign for $\gamma_2$. In the mirror case where $y$ and $\theta$ are both “bad” the expected sign for $\gamma_2$ would still be positive.

We cannot, however, rule out the possibility of an opposite sign for $\gamma_2$ because we do not observe $y$ and $\theta$ separately but only their weighted average $f^e$. A high $f^e$ could then be the result of a very high $y$ and a relatively low $\theta$, or vice versa. In these situations, the expected sign will depend on which of the two parameters is higher and on whether it is the precision of private or public information that is changing. For instance, the expected sign for $\gamma_2$ would still be positive if a high $f^e$ were due to a particularly high $y$ and the precision of public information $\alpha$ were changing; it would also be positive if a high $f^e$ were due to a particularly high $\theta$ and the precision of private information $\beta$ were changing. However, opposite combinations would yield the opposite expected sign.

In the theoretical model the probability of a speculative attack also depends on the potential gains in the event of a successful attack, namely $D - t$. One can show that an increase in the potential gross profit $D$ makes

\textsuperscript{16}Annex 1 shows that the variance of public information has similar effects in a model with only public information, independently of the number of equilibria. However, a proper test of a model with multiple equilibria would require a different econometric approach, one allowing for jumps across multiple equilibria.
an attack more likely. As an indicator of potential gross profits we select the real effective exchange rate \( e \). A rise in \( e \), i.e. a decrease in external competitiveness, may signal to speculators that the devaluation of the currency will be greater when the exchange rate regime collapses. Thus, if transaction costs \( t \) are constant, a rise in \( e \) may represent an increase in speculators’ potential gains. Since speculative pressures are increasing in potential gains, we expect \( \gamma_3 \) to be positive.

4 Empirical evidence

In this section, we verify whether agents’ expectations on economic fundamentals help explain actual exchange rate pressures. For this purpose, we build a monthly dataset with indices of exchange rate pressure and means and variances of Consensus Economics forecasts of GDP growth for six Asian countries (Thailand, South Korea, Indonesia, Malaysia, Singapore, and Hong Kong) from January 1995 to May 2001.

4.1 The data

To measure the fraction of speculators that decide to attack the currency, we build an index of exchange rate pressure.\(^{17}\) In recent years, several empirical studies have developed indicators of exchange rate pressure designed to identify and predict crisis periods. In this paper, we follow a similar methodology, except that we do not transform the index of exchange rate pressure into a discrete zero-one variable separating tranquil from crisis periods.\(^{18}\) The reason is that in practice some speculators attack the currency while others do not, consistently with the prediction of a private information model in which the number of speculators attacking a currency varies continuously with fundamentals and the distribution of beliefs.

Our index of exchange rate pressure \( \text{IND3} \) is the sum of the normalized values of three indicators of exchange rate pressure:\(^{19}\) i) the percentage depreciation of the domestic currency against the U.S. dollar over the previous month; ii) the fall in international reserves over the previous month as a percentage of the 12-month moving average of imports; and iii) the three-month interest rate less the annualized percentage change in consumer prices over the previous six months. To check the robustness of our results,\(^ {15}\)

\(^{17}\)Girton and Roper (1977), Roper and Turnovsky (1980), and Weymark (1998) discuss the assumptions needed to justify different definitions of indices of exchange rate pressure in theoretical macro models.

\(^{18}\)Another exception is Sachs et al. (1996) who use a weighted sum of the percentage decrease in reserves and the percentage depreciation of the exchange rate in a cross-country regression.

\(^{19}\)To normalize, we subtract from each indicator the country-specific mean and divide the result by the country-specific standard deviation.
we also compute an index $IND_2$, which sums only normalized values of $i)$ and $ii)$, and an index $BIS$, which is the continuous version of an index recently developed by the Bank for International Settlements for monitoring purposes. Figure 6 shows the time-series behavior of these three indices.

Every month, Consensus Economics gathers forecasts of a series of macro variables for the current and the following year. The panel of forecasters encompasses most international big players in the foreign exchange market as well as domestic enterprises; the former group includes both large banks such as Citigroup, HSBC, JP Morgan Chase, Barclays, UBS, ING Bank, ABN AMRO and important non-banking financial institutions such as Goldman Sachs, Lehman Brothers, Merrill Lynch and Morgan Stanley. The number of forecasters surveyed by Consensus Economics varies across countries and over time; its average for each of the six countries in our sample ranges between 14 and 18 and was slightly larger (between 15 and 20) in 1997.

Following Brooks et al. (2001), in order to reproduce a constant forecast horizon of one year, we compute a weighted average of current-year and following-year forecasts with weights equal respectively to 11/12 and 1/12 in January, 10/12 and 2/12 in February, and so on until 0/12 and 12/12 in December. To reduce the effect of possible outliers, we use the median (rather than the mean) of Consensus Economics forecasts at each date and the mean absolute median difference as a measure of dispersion. We limit our analysis to the forecasts of GDP growth. Consensus Economics forecasts for other variables—inflation, current account balance, trade balance, and exports—are available, but the number of forecasts is generally smaller.

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20 Indices based only on exchange rate and reserve changes are the most common in empirical works on early warning systems, because of the lack of reliable data on interest rates for panel datasets with a large number of developing countries and a long time series dimension. This is the case of the early warning system used by the IMF (see Berg et al. (2000)).

21 The BIS index is based on four indicators of exchange rate pressure: $i)$ the percentage depreciation of the domestic currency against the U.S. dollar over three months; $ii)$ the percentage depreciation of the domestic currency against the U.S. dollar over one year; $iii)$ the three-month interest rate less the annualized percentage change in consumer prices over the previous six months; and $iv)$ the fall in international reserves over three months as a percentage of the 12-month moving average of imports. The BIS transforms the values of each indicator into scores that are then weighted to compute an index that can take 21 different values from -10 (maximum appreciating pressure) to +10 (maximum depreciating pressure). Annex B of Hawkins and Klau (2000) describes the construction of this index in detail. By contrast, we compute a continuous index by adding normalized values of each of the four indicators of exchange rate pressure.

22 Multicollinearity of current-year and following-year forecasts prevents us from including both variables in the regression. However, very similar results were obtained by including only the following-year forecast, only the current-year forecast, or the following-year forecast together with the difference between the two. In these cases, the dispersion measures were seasonally adjusted to account for the smaller dispersion of forecasts—documented by Loungani (2001)—at the end of the year than at the beginning of the year.
than for GDP growth, making mean and dispersion measures less reliable. Moreover, in preliminary estimates, these other variables did not perform as well as GDP growth and, when measures of the mean and variance of expected GDP growth were included in the regression, hardly any other forecast variable was significant.

The real effective exchange rate is computed by JP Morgan and is generally available with a one-month lag. We found that the overall fit using the real effective exchange rate was marginally better than using the nominal exchange rate with the US dollar, but there was no difference in terms of the estimated signs and significance of all other coefficients between the two models. The actual values of GDP growth and other variables used in previous studies – such as inflation, international reserves, the ratio of M2 to international reserves, and the ratio of BIS external short-term debt to international reserves – had little effect on exchange rate pressures once we included the mean and variance of expected GDP growth in the regression.

4.2 Benchmark regression

Our benchmark regression is the following estimated version of equation (6):

\[
IND3_{j,t} = \hat{\gamma}_0 + \hat{\gamma}_1 f^{e}_{\text{GDP},j,t-1} + \hat{\gamma}_2 \sigma_{\text{GDP},j,t-1} \cdot (f^{e}_{\text{GDP},j,t-1} - \hat{\gamma}_3 \tilde{f}_{\text{GDP}}) \\
+ \hat{\gamma}_3 e_{j,t-1} + u_{j,t}, \quad u_{j,t} = \hat{\rho}_j u_{j,t-1} + \epsilon_{j,t}
\]  

(7)

where \(IND3_{j,t}\) is our three-component index of exchange rate pressure for country \(j\) at time \(t\). First, we estimated this system as a set of seemingly unrelated regressions (SUR) with country-specific coefficients and a country-specific AR(1) term to correct for serial correlation. We chose the SUR estimation method to allow for the likely correlation of the errors across countries during the Asian crisis. Second, we performed a Wald test of equality of parameters across countries, which showed that the coefficients \(\hat{\gamma}_1\) and \(\hat{\gamma}_2\) could be constrained to be the same across countries (the null hypothesis of equality was accepted with a p-value of 0.745). Table 1 shows the results of this restricted estimation of (7). We use the restrictions accepted by the data to simplify the presentation and to conduct robustness tests involving recursive estimation (see below) on a specification with a limited number of parameters. The restriction is by no means necessary to obtain statistically significant coefficients. In the unrestricted estimates, all \(\hat{\gamma}_{1,j}\) \((j = 1, \ldots, 6)\) were negative and statistically significant at the 5 percent confidence level and all \(\hat{\gamma}_{2,j}\) \((j = 1, \ldots, 6)\) were positive and statistically significant at the 1 percent confidence level.

The results in Table 1 confirm that higher expected GDP growth reduces exchange rate pressures. Most interestingly, these estimates indicate that uncertainty about GDP growth has an additional effect, which depends on...
the expected GDP growth, as our theoretical model predicts. A higher dispersion of GDP growth forecasts tends to increase exchange rate pressures when expected GDP growth is above the estimated country-specific threshold and to reduce them when it is below. (The threshold is statistically different from zero only for Singapore.) These findings are consistent with the theoretical result that a reduction in the precision of public and private information increases the number of speculators attacking the currency when the public signal and the average private signal are both “good” and has an opposite effect when they are both “bad” (Proposition 2).

Uncertainty about GDP growth is the most important variable in our regression and provides the largest contribution to the overall goodness-of-fit measures only when interacted with the mean of individual forecasts. The contribution of the interaction term to the overall goodness-of-fit measures is larger than that of the mean forecast alone: if we exclude $\sigma^e_{\text{GDP},t-1}$ by setting $\hat{\gamma}_1 = 0$ in (7) the $R^2$ for the overall system falls from 42.6 to 37.6 percent, whereas if we exclude $\sigma^e_{\text{GDP},t-1} \cdot (\bar{f}^e_{\text{GDP},t-1} - \hat{\gamma}_j \hat{\sigma}_{\text{GDP}})$ the $R^2$ for the overall system drops from 42.6 to 33 percent. Adding back the dispersion of the forecasts $\sigma^e_{\text{GDP},t-1}$ to the list of regressors yields an $R^2$ of only 34.3 percent. This is an additional confirmation of the model prediction that the effect of uncertainty depends critically on the level of expected fundamentals.

4.3 Sensitivity analysis
This section presents a series of robustness tests of our benchmark specification (7), confirming our main result that uncertainty about fundamentals plays a role in currency crises and that this role depends on expected fundamentals.

Robustness to alternative exchange rate pressure measures. Tables 2 and 3 present estimates of the specification (7) with two alternative measures of exchange rate pressure as dependent variable ($\text{IND2}$ and $\text{BIS}$). The coefficient measuring the effect of uncertainty ($\hat{\gamma}_2$) remains positive and strongly significant. The coefficient for the effect of expected fundamentals ($\hat{\gamma}_1$) remains negative and significant.

Robustness to fixed and floating exchange rate regimes. Should we test the implications of our model only on the pre-crisis sample? Prima facie, this approach would be consistent with the model of Section 2.2, in which the government pegs the exchange rate. Yet, there are theoretical and empirical reasons why the predictions of this model should be tested on the entire sample. From a theoretical point of view, in a floating exchange rate regime speculators still face a coordination problem: the future value of the currency and, in turn, their potential profits depend on how many
buy or sell the currency. Thus, each speculator still plays a coordination game with the others that might result in a tripartition of the space of fundamentals similar to that of second-generation currency crisis models. Assume, for instance, there are values of the fundamentals that are so good that an appreciation is certain, values that are so bad that a depreciation is certain, and values (maybe most values) for which the outcome depends on how many speculators decide to buy or sell. Within this model, the mean and variance of speculators’ expectations will produce downward or upward pressures on the currency similar to those we have obtained in Section 2.2.

From an empirical point of view, some countries in the sample – Hong Kong and Singapore – never changed their exchange rate regime, while Malaysia repegged its currency in September 1998. Moreover, the post-crisis regime of the other countries was not a free float but a managed float, whose features can still be captured by the model of Section 2.2. The countries that abandoned pegs, in fact, recorded the largest outflows of international reserves in the second half of 1997. As a result, for all these countries but South Korea, the greatest drop in reserves came after the change in the exchange rate regime. In some cases, the depletion of official reserves continued in the first quarter of 1998 and recurred after the Russian crisis.

These considerations suggest that the full-sample estimates of Tables 1, 2, and 3 represent a meaningful test of the model, but it is still interesting to verify whether our results would be changed by restricting the analysis to the pre-crisis period 1995:01-1997:07. Table 4 shows the outcome of this exercise. Because of the substantial reduction in the number of observations, we now also restrict $\gamma_1, \gamma_3, \text{ and } \hat{p}_j$ to be the same in all countries, allowing only the intercepts $\gamma_{0,j}$ in each equation to be country-specific. This is equivalent to estimating a panel model with fixed effects. The effect of uncertainty is positive and statistically significant in this pre-crisis period as well. The negative effect of better expected fundamentals on exchange rate pressures is also confirmed. These results hold with all measures of exchange rate pressure and confirm that the breakdown of the exchange rate regime in most of the countries in our panel in the second half of 1997 is not the sole cause of the estimated effect of uncertainty on exchange rate pressures.

This is consistent with the increase in uncertainty about GDP growth prior to the crisis in Thailand (from mid-1996), South Korea (from end-1996), and, to a lesser extent, Malaysia (Figure 2). The increase of uncertainty in Hong Kong – which maintained its currency board for the entire period – provides further evidence that the breakdown of the exchange rate regime may not be the only cause of the uncertainty we observe.

We further checked the robustness of our results by re-estimating the
benchmark SUR model of Table 1 with a set of dummies that were set to 1 when a country no longer pegged its exchange rate. The results were essentially unchanged, $\hat{\gamma}_1$ and $\hat{\gamma}_2$ remaining very significant. Nor did the results change when the model in Table 4 was estimated on unbalanced panels excluding either the observations following the breakdown of each country's exchange rate regime or the observations following each country’s maximum currency depreciation. Finally, the statistical significance of the pre-crisis recursive estimates of $\hat{\gamma}_1$ and $\hat{\gamma}_2$ (Figure 5) provides another indication that our results also hold in the pre-crisis sample.

Robustness to dynamic specification and endogenous regressors. As it is common in the literature, we exploit time-series information to test an essentially static game, implicitly assuming that the data come from a myopic, repeated play of the one-shot game. To be consistent with this approach, we correct for serial correlation of the errors by including a country-specific AR(1) term in our benchmark regression (7) rather than estimating a dynamic specification. In a possible dynamic extension of our theoretical model, speculators would use information revealed in previous stages of the game to decide whether to attack the currency in the current period. While developing a dynamic version of the model in Section 2 is beyond the scope of this paper, we can estimate a dynamic version of equation (7) with the lagged exchange rate pressure index on the right-hand side, implicitly assuming that the information on exchange rate pressures in the previous month is available to speculators at the beginning of the following month. Estimating this specification with SUR methodology yields results very similar to those reported in Table 1 confirming sign and statistical significance of all coefficients.

To correct for the possible bias due to the inclusion of the lagged dependent variable among the regressors, we also estimate a dynamic panel version of our model using the GMM estimator of Arellano and Bond (1991) with additional lags of the dependent variable as instruments. The Sargan’s test of overidentifying restrictions and Arellano and Bond’s test for (second order) residual autocorrelation confirm the validity of these instruments. The GMM estimates are in line with previous results: $\hat{\gamma}_1$ and $\hat{\gamma}_2$ have the same signs as in the benchmark regression and are both statistically significant at the 1 percent confidence level.

The GMM estimator allows us to check also the robustness of our results to the potential endogeneity of the mean and variance of the forecasts. Lagging the mean and variance of the forecasts by one month – as we do in the benchmark regression – rules out contemporaneous reverse causality. Predetermined mean and variance of the forecasts could, however, still be endogenous if a serially correlated omitted variable affected both exchange rate pressures and the distribution of the forecasts. In this case, lagged
values of the mean and variance of the forecasts and the serially-correlated error term would not be independent. We address this potential problem by instrumenting the one-period-lagged mean and variance of the forecasts with additional lags of the same variables and verifying the validity of these instruments with tests of overidentifying restrictions and residual autocorrelation. The GMM estimates easily pass these tests and confirm the signs and significance of $\hat{\gamma}_1$ and $\hat{\gamma}_2$ obtained with the benchmark specification and the SUR estimation technique. Allowing for country-specific coefficients in the GMM estimates does not change these results.

**Time-varying $\hat{\gamma}_1$ and $\hat{\gamma}_2$.** Another robustness check regards the possible instability over time of the coefficients $\hat{\gamma}_1$ and $\hat{\gamma}_2$. Proposition 2 implies that the effect of expected fundamentals on exchange rate pressures is always negative but may vary over time together with the precision of public and private information. We allow for this possibility by estimating $\hat{\gamma}_1$ recursively with state-space techniques. Figure 7 (top panel) shows that $\hat{\gamma}_1$ varies within a relatively narrow range, remaining always negative and strongly significant. Similarly, the effect of uncertainty on exchange rate pressures may vary depending not only on the level of expected fundamentals (for which we control) but also on whether it is the precision of public or private information that changes and on the difference between the actual fundamental $\theta$ and the cutoff point $x^*$. In particular, there may be instances in which changes in the precision of private information may cause the parameter $\hat{\gamma}_2$ to turn negative. We check this possibility by estimating $\hat{\gamma}_2$ recursively. Figure 7 (bottom panel) shows that the recursive estimated $\hat{\gamma}_2$ changes over time but remains always positive and significantly different from zero.\(^25\)

**Time-varying threshold $\hat{\gamma}$.** The last robustness check is the estimation of the thresholds separating “high” from “low” expected GDP growth. These are also likely to be time-varying, reflecting changes in the parameters in $s_1$, $s_2$, and $s_3$ or, more simply, because investors might have revised estimates of potential growth rates as the crisis progressed. To address this potential concern, we estimate the six parameters $\hat{\gamma}_j$ in (7) recursively (Figure 8). In all countries except Hong Kong, the estimated thresholds tend to decline until end-1997 before rebounding and stabilizing below their pre-crisis level. Nevertheless, Table 5 shows that allowing for time-varying thresholds has little effect on $\hat{\gamma}_1$ and $\hat{\gamma}_2$; the latter remains strongly significant and positive. Note that the overall estimated effect of $\sigma_{GDP_j,t-1}^2$

\(^{25}\)We also estimated separate recursive coefficients $\hat{\gamma}_{2,j}$ for each country. Because of the smaller number of observations, the country-specific estimates had larger RMSE bands than those in Figure 7 at the beginning of the period. The estimated coefficients were, however, mostly positive with a statistically significant negative coefficient only for the early part of the Hong Kong sample.
on exchange rate pressures (measured by $\gamma_2 \cdot \left( f_{GDP,t-1}^e - \hat{\gamma}_{GDP,t-1} \right)$) may also vary with changes in GDP forecasts ($f_{GDP,t-1}^e$) and country-specific thresholds ($\hat{\gamma}_{GDP,t-1}$). Figure 9 shows that this estimated effect varies substantially over time but remains mostly positive, with the exception of Indonesia in 1998-99 and Singapore at end-1998.

5 Conclusions

This paper studies how uncertainty about fundamentals contributes to currency crises, both theoretically and empirically. While there are several challenges in developing an empirical test of Morris and Shin’s model of currency crises, we are able to relate precisely the mean and the variance of Consensus Economic forecasts to the parameters of the model. This allows us to test whether: i) improvements in the public signal and in the mean of the private signals are associated with lower exchange rate pressures; ii) a greater precision of public and private information has an effect on exchange rate pressures that is additional to that of the public and private signals; and iii) the effect of information precisions depends on how “good” the public and the private signals are.

Our estimates on a monthly dataset of forecasts for six Asian countries confirm that both the mean and the variance of agents’ expectations about economic fundamentals contribute to explaining exchange rate pressures. Specifically, exchange rate pressures diminish with an improvement in the mean forecast of GDP growth, and increase with the dispersion of GDP growth forecasts when expected growth is relatively high.

Estimates of the threshold separating good from bad expected GDP growth imply that in all the countries in our sample uncertainty about GDP growth increased exchange rate pressures in the pre-crisis period (before July 1997) and after mid-1999 (Figure 9). During the intermediate period, in some countries uncertainty about the growth outlook had a significant attenuating effect on exchange rate pressures. This effect was temporary and was greatest at the time of the Russian crisis (end-1998), which coincided with a period of low expected growth.

These results are robust to the definition of exchange rate pressure indices and to the location of the threshold separating good from bad growth outlooks. Moreover, the significant role of uncertainty even in the pre-crisis period alone implies that the collapse of the exchange rate regime in most countries in the sample is not the sole determinant of our results.

How does our paper relate to previous explanations of the Asian crisis? While our focus on the role of uncertainty about fundamentals is unique, our theoretical model is not in contrast with the prevailing views of the Asian crisis. Previous theoretical studies have interpreted the crisis as a jump from a “good” to a “bad” equilibrium due to the presence of internationally

22
illiquid banks (see Chang and Velasco (2001) and Morris and Shin (2004)) or credit constraints (Aghion et al. (2001)). Our stylized model with imprecise public information, but no private information, also yields multiple equilibria (Annex 1). In the public information version of the model, the mean and variance of public information have effects that are qualitatively similar to those in the model with both public and private information. Moreover, with both kinds of information (just as in multiple equilibria models), if speculators coordinate crises can take place for values of the fundamentals that – without an attack – would be consistent with a sustained peg. The main difference from multiple equilibria models is that sufficiently precise private information yields a unique equilibrium that allows us to determine exactly the level of fundamentals and precisions of information for which there are enough speculators that coordinate and bring down the peg.

Our empirical results are also broadly consistent with previous explanations of the Asian crisis based on multiple equilibria models. A deterioration of expectations about the viability of the Asian corporate and banking sectors is, in fact, consistent with the significant and large impact of expectations about future GDP growth (and their dispersion) that we find in our estimates.

While a welfare analysis of the provision of public information is beyond the scope of this paper, our results do shed light on whether a country may better resist a speculative attack on its currency when the precision of the official information it releases is high. Both theoretical and empirical results suggest that the precision of public information may either help or hurt a country under attack, depending on the state of fundamentals. The theoretical model predicts that the precision of public information helps when the public signal is good, but hurts when it is bad. Unsurprisingly, transparent policies may then benefit “virtuous” countries. The empirical results suggest that at the onset of the Asian crisis, when expected fundamentals were still relatively good but uncertainty was increasing, a higher precision of official information would have been beneficial. At the same time, there is some indication that during some phases of the crisis uncertainty about the economic outlook may have dampened speculative pressures. However, appropriate discussion of the welfare implications of the precision of official information would require developing a theoretical model in which speculators factor the authorities’ strategy of information releasing into their decisions.

Future theoretical research could also verify whether the effect of the precisions of public and private information on the share of speculators who decide to attack the currency is robust to the choice of the payoff function and the probability distribution. Relaxing the assumption of exogenous fundamentals and exploring the consequences of exchange rate changes that have feedback effects on economic fundamentals could also have interesting implications.
Future empirical research is also needed to verify whether data on other well-known currency crises in Latin America and Europe confirm the statistical significance of uncertainty about fundamentals. There may also be scope for an empirical verification of the multiple equilibria model with regime switching econometric techniques as in Jeanne (1997) and Jeanne and Masson (2000). While testing the leading indicator properties of the mean and variance of Consensus Economics forecasts is beyond the scope of this paper, it would be worthwhile exploring whether these variables can enhance the predictive power of early warning systems, which are currently based only on past fundamentals. In this regard, the results of our estimates on the pre-crisis period are promising.
A Appendix

A.1 Equilibrium of the public and private information game

In this section we characterize the unique equilibrium of the game with both public and private information. To provide an intuition for the mechanism leading to a unique equilibrium, we can use the infection argument, as in Morris et al. (1995). Suppose that a speculator is known to undertake a certain action given some private information set. This knowledge might imply a unique best response by the other speculators given some of their information sets where the first information set is considered possible. This, in turn, may imply that the original speculator responds to that knowledge by choosing that same action on a larger information set, and so on. In the currency crisis game, if private information is sufficiently precise, this chain of reasoning results in a unique action profile, eliciting a unique equilibrium.

We now turn to the problem of characterizing the equilibrium. Morris and Shin (1998 and 2004) and Metz (2002) have shown that the unique equilibrium can be specified by a couple $(x^*, \theta^*)$, such that speculators use the trigger strategy

$$
\delta(x) = \begin{cases} 
\text{attack} & \text{if } x \leq x^* \\
\text{don't attack} & \text{if } x > x^* 
\end{cases},
$$

and the government follows the rule:\footnote{Given $\theta$, the share of attackers is completely determined by $\delta$, since we have assumed that there is a continuum of speculators. It follows that when speculators use $\delta$, the function $\psi$ below is exactly the same as the government’s decision rule specified in Section 2 (which was therefore denoted by the same symbol $\psi$).}

$$
\psi(\theta) = \begin{cases} 
\text{abandon} & \text{if } \theta \leq \theta^* \\
\text{defend} & \text{if } \theta > \theta^* 
\end{cases}.
$$

Here, we first assume that agents use a trigger strategy like $\delta$; we then derive a sufficient condition granting that unique values of $x^*$ and $\theta^*$ exist; finally, we find the equations that characterize these values. We do not show that a trigger strategy for speculators is the unique optimal strategy under the sufficient condition, as this result follows directly from Morris and Shin (2004) or, in a more general framework, from Frankel et al. (2003).

Trigger points

Assume that agents use the trigger strategy $\delta$ defined above and let us find the trigger point of the government’s optimal strategy. Given $x^*$ and $\theta$, the share of speculators attacking the currency is

$$
\Pr(X < x^* \mid \theta) = \Phi \left( \sqrt{3} (x^* - \theta) \right).
$$
As the expected utility from abandoning the peg is nil, the government is indifferent between defending and abandoning the peg for the level of fundamentals $\theta^*$ that solves:

$$\theta^* - \Phi \left[ \sqrt{\beta} (x^* - \theta^*) \right] = 0 .$$  \hspace{1cm} (8)

Equation (8) implicitly defines $\theta^*$ as a function of $x^*$. Note that $\Phi$ is decreasing and continuous in $\theta^*$, and takes all the values in the open interval $(0, 1)$. Therefore, there exists a unique value of $\theta^*$ that solves (8), for any $x^* \in \mathbb{R}$.

Let us find the trigger point for speculators. Given $\psi$, the expected utility of a speculator who receives a message $x$ and attacks the currency is:

$$(D - t) \cdot \Pr(\Theta \leq \theta^* \mid x) - t \cdot \Pr(\Theta > \theta^* \mid x) = D \cdot \Pr(\Theta \leq \theta^* \mid x) - t .$$

As the expected utility from don’t attack is nil, a speculator is indifferent between attacking and not when he receives the message $x^*$ that solves:

$$D \cdot \Phi \left[ \sqrt{\alpha + \beta} \left( \theta^* - \frac{\alpha}{\alpha + \beta} y - \frac{\beta}{\alpha + \beta} x^* \right) \right] - t = 0 .$$  \hspace{1cm} (9)

**Sufficient condition for a unique equilibrium**

Unlike equation (8), equation (9) does not necessarily have a unique solution. Note that, as $x^*$ goes to $-\infty$, the left-hand side of equation (9) goes to $D - t > 0$; when $x^*$ goes to $+\infty$, the left-hand side of equation (9) goes to $-t < 0$. By the continuity of the left-hand side of (9), a sufficient condition granting that a unique solution to equation (9) exists may be obtained by requiring that the derivative of the left-hand side of (9) with respect to $x^*$ is smaller than zero; namely:

$$D \cdot \sqrt{\alpha + \beta} \left( \frac{d\theta^*}{dx^*} - \frac{\beta}{\alpha + \beta} \right) \cdot \phi \left[ \sqrt{\alpha + \beta} \left( \theta^* - \frac{\alpha}{\alpha + \beta} y - \frac{\beta}{\alpha + \beta} x^* \right) \right] < 0 .$$

The previous inequality holds provided that

$$\frac{d\theta^*}{dx^*} - \frac{\beta}{\alpha + \beta} < 0 .$$  \hspace{1cm} (10)

Differentiating implicitly equation (8) we can obtain

$$\frac{d\theta^*}{dx^*} = \frac{\sqrt{\beta} \cdot \phi \left[ \sqrt{\beta} (x^* - \theta^*) \right]}{1 + \sqrt{\beta} \cdot \phi \left[ \sqrt{\beta} (x^* - \theta^*) \right]} ,$$

and, substituting into (10),

$$\sqrt{\beta} \frac{1}{\phi\left[ \sqrt{\beta} (x^* - \theta^*) \right]} + \sqrt{\beta} < \frac{\beta}{\alpha + \beta} .$$  \hspace{1cm} (11)
A sufficient condition for inequality (11) to hold is:

\[ \frac{\sqrt{\beta}}{\frac{1}{\max x\phi(x)} + \sqrt{\beta}} < \frac{\beta}{\alpha + \beta}. \]

Rearranging the previous inequality – and recalling that the maximum of \( \phi \) is \( 1/\sqrt{2\pi} \) – we obtain the sufficient condition (1).

**Equilibrium**

Given the sufficient condition (1), the unique equilibrium is characterized by \( (x^*, \theta^*) \) which are determined by the unique solution of the following system of equations:

\[
\begin{align*}
0 &= \theta^* - \Phi \left[ \sqrt{\beta} \left( x^* - \theta^* \right) \right] \\
0 &= D \cdot \Phi \left[ \sqrt{\alpha + \beta} \left( \theta^* - \frac{\alpha}{\alpha + \beta} y - \frac{\beta}{\alpha + \beta} x^* \right) \right] - t.
\end{align*}
\]  

(12)

**A.2 Expectation effects on the trigger points**

We now show that both \( \theta^* \) and \( x^* \) are decreasing in \( y \) and that the effect of the precision of public information depends on the ex-ante expected fundamental: if \( y \) is sufficiently good (bad), then an increase in \( \alpha \) makes \( \theta^* \) and \( x^* \) decrease (increase). Moreover, we prove that an increase in the precision of private information \( \beta \) has the reverse effect, making \( \theta^* \) and \( x^* \) increase (decrease) when \( y \) is sufficiently good (bad).

The three conditions for \( \alpha \) to reduce \( \theta^* \) and \( x^* \), for \( \beta \) to raise \( \theta^* \), and for \( \beta \) to raise \( x^* \) respectively are:\[27]

\[
\begin{align*}
y &> \theta^* - \frac{1}{2\sqrt{\alpha + \beta}} \Phi^{-1} \left( \frac{t}{D} \right) \equiv s_1 \\
y &> \theta^* - \frac{1}{\sqrt{\alpha + \beta}} \Phi^{-1} \left( \frac{t}{D} \right) \equiv s_2 \\
y &> \theta^* - \frac{\alpha^2 \phi - 2\sqrt{\beta} \alpha - (\sqrt{\beta})^3}{\alpha \sqrt{(\alpha + \beta)(\alpha \phi - \beta \phi - 2\sqrt{\beta})}} \Phi^{-1} \left( \frac{t}{D} \right) \equiv s_3.
\end{align*}
\]  

More precisely, the effects of expectations on the trigger point of the government’s strategy, \( \theta^* \), can be summarized by the following result by Metz (2002):

**Proposition 3 (Metz, 2002)** Assume that \( \beta > \frac{2}{2\pi} \). Then \( \theta^* \) is: (i) decreasing in \( y \); (ii) decreasing (increasing) in \( \alpha \) if \( y > s_1 \) \( (y < s_1) \); (iii) increasing (decreasing) in \( \beta \) if \( y > s_2 \) \( (y < s_2) \).

\[27\] Note that, if \( D = 2t \), the conditions (13), (14), and (15) coincide.
The effects of the parameters on the decision rule of speculators (i.e. on the trigger point $x^*$), which are crucial to derive the results on the share of attackers presented in Proposition 2, are given by the following proposition:

**Proposition 4** Assume that $\beta > \frac{\alpha^2}{2\pi}$. Then $x^*$ is: (i) decreasing in $y$; (ii) decreasing (increasing) in $\alpha$ if $y > s_1$ ($y < s_1$); (iii) increasing (decreasing) in $\beta$ if $y > s_3$ ($y < s_3$).

The effects of the parameters on $\theta^*$ and $x^*$ are essentially the same. An increase in $y$, by reducing $\theta^*$ and $x^*$, makes a currency crisis less likely. Propositions 3 and 4 also show that the effect of the precision of the public signal $\bar{\beta}$ depends on $y$. If $\bar{\beta}$ increases and $y$ is sufficiently good (bad), $\theta^*$ and $x^*$ decrease and a speculative attack becomes less (more) likely.

In order to derive the effects of the parameters $y$, $\alpha$, $\beta$ on $(x^*, \theta^*)$ the system (12) can also be written as

$$
\begin{align*}
x^* &= \theta^* + \frac{1}{\sqrt{\beta}} \Phi^{-1} (\theta^*) \\
x^* &= \frac{\alpha + \beta}{\beta} \theta^* - \frac{\alpha}{\beta} y - \frac{\sqrt{\alpha + \beta}}{\beta} \Phi^{-1} \left( \frac{t}{D} \right) 
\end{align*}
$$

that, by substitution, yields:

$$
\theta^* = \Phi \left[ \frac{\alpha}{\sqrt{\beta}} \left( \theta^* - y - \frac{\sqrt{\alpha + \beta}}{\alpha} \Phi^{-1} \left( \frac{t}{D} \right) \right) \right].
$$

In the following, by differentiating the system of implicit equations (12) (or the alternative expressions (16)) we simultaneously obtain the effect of each parameter on both $\theta^*$ and $x^*$, thereby proving both propositions 3 and 4.

**Effects of $y$**

By differentiating the system of implicit equations (12) with respect to $y$, we can obtain:

$$
\begin{align*}
0 &= \frac{d\theta^*}{dy} - \frac{\sqrt{\beta} \left( dx^* - \frac{d\theta^*}{dy} \right) \phi}{\sqrt{\beta - \alpha \phi}} \\
0 &= \frac{d\theta^*}{dy} - \frac{\alpha}{\alpha + \beta} - \frac{\beta}{\alpha + \beta} \frac{dx^*}{dy},
\end{align*}
$$

where we have neglected the argument of $\phi$. Solving by substitution, we get:

$$
\begin{align*}
\frac{d\theta^*}{dy} &= -\frac{\alpha \phi}{\sqrt{\beta - \alpha \phi}} \\
\frac{dx^*}{dy} &= -\frac{\alpha}{\beta} + \frac{\alpha + \beta}{\beta} \frac{d\theta^*}{dy}.
\end{align*}
$$
Therefore, the derivative of $\theta^*$ with respect to $y$ is negative, provided that $\beta > \alpha^2 \phi^2$. But this inequality certainly holds under the sufficient condition (1). Hence, $d\theta^*/dy$ is negative and, in turn, $dx^*/dy$ is negative too.

Effects of $\alpha$
In order to derive the effect of $\alpha$ on $\theta^*$, we can simplify our calculations starting by differentiating equation (17):

$$
\frac{d\theta^*}{d\alpha} = \left(\frac{\theta^*}{\sqrt{\beta}} + \frac{\alpha}{\sqrt{\beta}} \frac{d\theta^*}{d\alpha} - \frac{y}{\sqrt{\beta}} - \frac{1}{2\sqrt{\beta}} \Phi^{-1}\left(\frac{t}{D}\right)\right) \cdot \phi
$$

where we have neglected the argument of $\Phi$. Solving for $d\theta^*/d\alpha$ we obtain:

$$
\frac{d\theta^*}{d\alpha} = \phi \cdot \left(1 - \frac{\alpha \phi}{\sqrt{\beta}}\right)^{-1} \cdot \left(\frac{\theta^*}{\sqrt{\beta}} - \frac{y}{\sqrt{\beta}} - \frac{1}{2\sqrt{\beta}} \Phi^{-1}\left(\frac{t}{D}\right)\right).
$$

The sufficient condition for a unique equilibrium (1) grants that the second term on the right-hand side of the previous equation is positive. By rearranging the third term, we find that the derivative of $\theta^*$ with respect to $\alpha$ is negative, provided that condition (13) holds.

Let us turn to the effect of $\alpha$ on $x^*$. Differentiating the first equation of system (12) with respect to $\alpha$ we get:

$$
\frac{d\theta^*}{d\alpha} - \left(\sqrt{\beta} \frac{dx^*}{d\alpha} - \sqrt{\beta} \frac{d\theta^*}{d\alpha}\right) \cdot \phi = 0,
$$

from which we can obtain:

$$
\frac{dx^*}{d\alpha} = \left(1 + \frac{1}{\phi \sqrt{\beta}}\right) \frac{d\theta^*}{d\alpha}.
$$

As the term in brackets is positive, the sign of the derivative of $x^*$ is the same as the sign of the derivative of $\theta^*$.

Effects of $\beta$ on $\theta^*$

Let us differentiate equation (17) with respect to $\beta$:

$$
\frac{d\theta^*}{d\beta} = \left(-\frac{\alpha}{2\sqrt{\beta^3}} \theta^* + \frac{\alpha}{\sqrt{\beta^3}} \frac{d\theta^*}{d\beta} + \frac{\alpha}{2\sqrt{\beta^3}} y + \frac{\alpha}{2\beta^2} \sqrt{\beta} \Phi^{-1}\left(\frac{t}{D}\right)\right) \cdot \phi,
$$

where we have neglected the argument of $\Phi$. Solving by substitution, we get:

$$
\frac{d\theta^*}{d\beta} = \phi \cdot \left(1 - \frac{\alpha \phi}{\sqrt{\beta}}\right)^{-1} \cdot \left(-\frac{\alpha}{2\sqrt{\beta^3}} \theta^* + \frac{\alpha}{2\sqrt{\beta^3}} y + \frac{\alpha}{2\beta^2} \sqrt{\beta} \Phi^{-1}\left(\frac{t}{D}\right)\right).
$$

29
The first two terms on the right-hand side of the previous equation are positive. By rearranging the third term, we get that the derivative of $\theta^*$ with respect to $\beta$ is positive, provided that condition (14) holds.

_Effect of $\beta$ on $x^*$_

Consider the second equation in system (16) and differentiate it with respect to $\beta$:

$$\frac{dx^*}{d\beta} = \frac{\alpha + \beta d\theta^*}{\beta} + \frac{\alpha (y - \theta^*)}{\beta^2} + \frac{2\alpha + \beta}{2\beta \sqrt{\alpha + \beta}} \Phi^{-1} \left( \frac{t}{D} \right).$$

Substituting the expression of $d\theta^*/d\beta$ previously found we can get – after some tedious algebra – that $dx^*/d\beta > 0$ iff

$$y > \theta^* - \frac{\alpha^2 \phi - 2\sqrt{3} \alpha - (\sqrt{3})^3}{\alpha \sqrt{\alpha + \beta} (\alpha \phi - \beta \phi - 2\sqrt{3})} \Phi^{-1} \left( \frac{t}{D} \right).$$
References


Figure 1: Mean and median forecasts of GDP growth
(weighted average of current and following year forecasts; 1995:01-2001:05)

Note: the shaded area marks the period from July 1997 to the end of 1998, which includes the Asian crisis, the Russian crisis, and the near-collapse of the hedge fund Long Term Capital Management.
Figure 2: Standard deviation and mean absolute median difference of GDP growth (weighted average of current and following year forecasts; 1995:01-2001:05)

Note: the shaded area marks the period from July 1997 to the end of 1998, which includes the Asian crisis, the Russian crisis, and the near-collapse of the hedge fund Long Term Capital Management.
Figure 3: Effects on share of attackers of higher $y$ and $\theta$.

Note: the thick line is the share of attackers (as a function of $\theta$) for $\alpha = \beta = t = 1$, $D = 2$, and $y = 0.5$ (which implies $\theta^* = x^* \simeq 0.500$). The dashed line is the share of attackers (as a function of $\theta$) for $\alpha = \beta = t = 1$, $D = 2$, and a "good" $y = 0.6$ (which implies $\theta^* = 0.434$ and $x^* \simeq 0.268$).
Figure 4: Effect on share of attackers of an increase in $\alpha$ for $y$ "good"

Note: the thick line is the share of attackers (as a function of $\theta$) for $\alpha = \beta = t = 1$, $D = 2$, and a "good" $y = 0.6$ (which implies $\theta^* = 0.434$ and $x^* \simeq 0.268$). The dashed line is the share of attackers (as a function of $\theta$) for $\beta = t = 1$, $D = 2$, $y = 0.6$, and $\alpha = 1.25$ (which implies $\theta^* = 0.402$ and $x^* \simeq 0.156$).
Figure 5: Effects on share of attackers of an increase in $\beta$ for $y$ “good”

Note: the thick line is the share of attackers (as a function of $\theta$) for $\alpha = \beta = t = 1$, $D = 2$, and $y = 0.6$ ($y$ is “good” since $x^* \simeq 0.268$). The thin line is the share of attackers for $\beta$ increased to 4 ($x^*$ raises to 0.444). The dotted line singles out the indirect effect of $\beta$ as it shows the share of attackers with the new $x^* = 0.444$ and the old $\beta = 1$. 


Figure 6: Indices of exchange rate pressure (1995:01 - 2001:05)

Note: the shaded area marks the period from July 1997 to the end of 1998, which includes the Asian crisis, the Russian crisis, and the near-collapse of the hedge fund Long Term Capital Management.
Figure 7: Recursive estimates of $\hat{\gamma}_1$ and $\hat{\gamma}_2$ (1995:08-2001:05)

Note: the shaded area marks the period from July 1997 to the end of 1998, which includes the Asian crisis, the Russian crisis, and the near-collapse of the hedge fund Long Term Capital Management.
Figure 8: Recursive estimates of the threshold separating high from low expected GDP growth (1996:07 - 2001:05)

Note: the shaded area marks the period from July 1997 to the end of 1998, which includes the Asian crisis, the Russian crisis, and the near-collapse of the hedge fund Long Term Capital Management.
Figure 9: Overall effect of uncertainty on exchange rate pressures in estimates with recursive threshold (1996:07-2001:05)

Thailand

Indonesia

Korea

Malaysia

Singapore

Hong Kong

Note: the shaded area marks the period from July 1997 to the end of 1998, which includes the Asian crisis, the Russian crisis, and the near-collapse of the hedge fund Long Term Capital Management.
### Table 1. Exchange Rate Pressure (IND3 Index) Estimates

(SUR estimates; standard errors in parenthesis; sample: 1995:03-2001:05 )

<table>
<thead>
<tr>
<th></th>
<th>Thailand</th>
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<tr>
<td>$\gamma_{0,j}$</td>
<td>-10.645 ***</td>
<td>-1.654</td>
<td>-17.254 ***</td>
<td>-13.830 ***</td>
<td>-18.052 ***</td>
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<td></td>
<td>(3.234)</td>
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<td>(3.439)</td>
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<td></td>
<td></td>
<td></td>
<td>-0.520 ***</td>
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<tr>
<td>$\gamma_2$</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td>0.592 ***</td>
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<td>$\rho_{j}$</td>
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<td>0.238 ***</td>
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<td>0.184 ***</td>
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<td>0.503 ***</td>
<td>0.345 ***</td>
<td>0.203 **</td>
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Data are monthly.
Three (***), two (**), and one (*) stars mark statistical significance respectively at one, five, and ten percent levels.
The coefficients $\gamma_1$ and $\gamma_2$ are restricted to be the same across countries.
Table 2. Exchange Rate Pressure (IND2 Index) Estimates
(SUR estimates; standard errors in parenthesis; sample: 1995:03-2001:05)

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<td>γ1</td>
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<td></td>
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<td>ρj</td>
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Observations | 76 | 76 | 76 | 76 | 76 | 76

Data are monthly.
Three (***), two (**), and one (*) stars mark statistical significance respectively at one, five, and ten percent levels.
The coefficients γ1 and γ2 are restricted to be the same across countries.
### Table 3. Exchange Rate Pressure (BIS Index) Estimates

(SUR estimates; standard errors in parenthesis; sample: 1995:03-2001:05)¹

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<td>( \rho_{1,j} )</td>
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<td>0.514 ***</td>
<td>0.510 ***</td>
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<td></td>
<td>(0.096)</td>
<td>(0.102)</td>
<td>(0.083)</td>
<td>(0.081)</td>
<td>(0.075)</td>
<td>(0.088)</td>
</tr>
<tr>
<td>( \rho_{2,j} )</td>
<td>-0.123 *</td>
<td>0.148 *</td>
<td>-0.452 ***</td>
<td>0.037</td>
<td>-0.255 ***</td>
<td>0.170 **</td>
</tr>
<tr>
<td></td>
<td>(0.072)</td>
<td>(0.081)</td>
<td>(0.069)</td>
<td>(0.074)</td>
<td>(0.071)</td>
<td>(0.078)</td>
</tr>
<tr>
<td>( R^2 )</td>
<td>0.258</td>
<td>0.452</td>
<td>0.599</td>
<td>0.323</td>
<td>0.297</td>
<td>0.353</td>
</tr>
<tr>
<td>DW</td>
<td>1.815</td>
<td>2.170</td>
<td>1.853</td>
<td>1.894</td>
<td>1.959</td>
<td>2.245</td>
</tr>
<tr>
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<td>76</td>
<td>76</td>
<td>76</td>
<td>76</td>
<td>76</td>
<td>76</td>
</tr>
</tbody>
</table>

¹ Data are monthly.

Three (***), two (**), and one (*) stars mark statistical significance respectively at one, five, and ten percent levels.
The coefficients \( \gamma_1 \) and \( \gamma_2 \) are restricted to be the same across countries.
<table>
<thead>
<tr>
<th></th>
<th>IND3</th>
<th>IND2</th>
<th>BIS</th>
</tr>
</thead>
<tbody>
<tr>
<td>$\gamma_1$</td>
<td>-1.450 *** (0.328)</td>
<td>-1.140 *** (0.227)</td>
<td>-1.632 *** (0.427)</td>
</tr>
<tr>
<td>$\gamma_2$</td>
<td>3.073 *** (0.862)</td>
<td>2.198 *** (0.642)</td>
<td>3.185 *** (1.098)</td>
</tr>
<tr>
<td>$\gamma$</td>
<td>7.297 *** (0.267)</td>
<td>7.572 *** (0.344)</td>
<td>6.996 *** (0.244)</td>
</tr>
<tr>
<td>$\gamma_3$</td>
<td>-0.008 (0.024)</td>
<td>-0.025 (0.018)</td>
<td>0.029 (0.029)</td>
</tr>
<tr>
<td>$\rho$</td>
<td>0.375 *** (0.066)</td>
<td>0.149 * (0.077)</td>
<td>0.608 *** (0.065)</td>
</tr>
<tr>
<td>$R^2$</td>
<td>0.384</td>
<td>0.094</td>
<td>0.392</td>
</tr>
<tr>
<td>DW</td>
<td>1.507</td>
<td>1.477</td>
<td>1.577</td>
</tr>
<tr>
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<td>174</td>
</tr>
</tbody>
</table>

1 Data are monthly. 
Three (***) , two (**) , and one (*) stars mark statistical significance respectively at one, five, and ten percent levels. 
The panel includes Thailand, Indonesia, South Korea, Malaysia, Singapore, Hong Kong.
Table 5. Exchange Rate Pressure (IND3 Index) Estimates with recursive threshold
(state-space estimates; standard errors in parenthesis; sample: 1995:03-2001:05)

<table>
<thead>
<tr>
<th></th>
<th>Thailand</th>
<th>Indonesia</th>
<th>South Korea</th>
<th>Malaysia</th>
<th>Singapore</th>
<th>Hong Kong</th>
</tr>
</thead>
<tbody>
<tr>
<td>$\gamma_{0,j}$</td>
<td>-9.609 (8.147)</td>
<td>-1.294 (1.604)</td>
<td>-12.448 (8.576)</td>
<td>-14.402 * (8.020)</td>
<td>-18.941 (12.549)</td>
<td>7.397 (6.090)</td>
</tr>
<tr>
<td>$\gamma_{1}$</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td>-0.290 * (0.159)</td>
<td></td>
</tr>
<tr>
<td>$\gamma_{2}$</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td>0.400 *** (0.085)</td>
<td></td>
</tr>
<tr>
<td>$\gamma_{j}$</td>
<td>estimated recursively (see Fig. 6)</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>$\gamma_{3,j}$</td>
<td>0.113 (0.087)</td>
<td>0.034 * (0.020)</td>
<td>0.168 * (0.102)</td>
<td>0.153 ** (0.071)</td>
<td>0.178 (0.113)</td>
<td>-0.062 (0.046)</td>
</tr>
<tr>
<td>$\rho_{j}$</td>
<td>0.385 ** (0.159)</td>
<td>0.250 (0.182)</td>
<td>0.487 * (0.264)</td>
<td>0.418 (0.273)</td>
<td>0.285 (0.209)</td>
<td>0.154 (0.207)</td>
</tr>
<tr>
<td>Observations</td>
<td>76</td>
<td>76</td>
<td>76</td>
<td>76</td>
<td>76</td>
<td>76</td>
</tr>
</tbody>
</table>

\[ Data\ are\ monthly.\ 
Three (***), two (**), and one (*) stars mark statistical significance respectively at one, five, and ten percent levels. 
The coefficients $\gamma_{1}$ and $\gamma_{2}$ are restricted to be the same across countries. \]
ANNEX:
A model with only public information
In this section we derive the effects of the ex-ante expected fundamental and of the precision of public information in a model with only public information. This model is relevant because it implies effects of these parameters similar to those of the unique-equilibrium model of Section 2.2, even though multiple equilibria are now possible. Specifically, also in this model, the way in which uncertainty contributes to currency crises depends on the ex-ante expected fundamental, thereby providing some further theoretical support to the empirical evidence of Section 4.

We assume that speculators have prior expectations about $\theta$ given by the same probability distribution $\Theta$ considered in Section 2.2; namely, $\Theta \sim \text{Norm}(y, 1/\alpha)$. The government observes both $\theta$ and $l$ before taking his decision, therefore his optimal strategy is the same function, $\psi$, as in the complete information model. Hence, if $\theta$ falls within $(-\infty, 0]$ the government devalues the currency, whilst if $\theta$ falls within $(1, +\infty)$ the government maintains the peg. When $\theta$ belongs to $(0, 1]$, speculators’ expectations will determine the outcome of the game.

Given $\psi$, we can focus on the Bayesian Nash equilibria of the reduced-form game of speculators. We need to calculate the expected payoff – denoted by $u(a_i, a_{-i})$ – of a speculator who attacks the currency when all other speculators also attack, and the expected payoff – denoted by $u(a_i, d_{-i})$ – of a speculator who attacks the currency when none do. Analytically these expected payoffs are given by:

$$u(a_i, a_{-i}) = \int_{-\infty}^{1} (D - t) \eta(\theta) d\theta - \int_{1}^{+\infty} t\eta(\theta) d\theta$$

$$u(a_i, d_{-i}) = \int_{-\infty}^{0} (D - t) \eta(\theta) d\theta - \int_{0}^{+\infty} t\eta(\theta) d\theta$$

where $\eta$ is the probability density function of $\Theta$. The following proposition specifies the Bayesian Nash equilibria of the reduced-form game of speculators:

**Proposition 1** The (“attack”) strategy profile in which all agents attack the currency is an equilibrium iff $u(a_i, a_{-i}) \geq 0$. The (“don’t-attack”) strategy profile in which all agents refrain from attacking is an equilibrium iff $u(a_i, d_{-i}) \leq 0$.

As $u(a_i, a_{-i})$ is always greater than or equal to $u(a_i, d_{-i})$, the “attack,” the “don’t-attack,” or both strategy profiles are equilibria of this game. Let

---

1In the unique-equilibrium model with both public and private information, comparative statics exercises predicted the likelihood of a speculative attack. In the model with only public information of this Appendix, which may yield multiple equilibria, we refer to a change in the likelihood of an attack more loosely – as it is common in the literature on speculative attacks – by relating it to the change in the range of parameters in which the attack strategy profile is an equilibrium.
us rewrite the two expected payoffs as:

\[
\begin{align*}
    u(a_i, a_{-i}) &= D \cdot \Phi \left[ \sqrt{\alpha} (1 - y) \right] - t \\
    u(a_i, d_{-i}) &= D \cdot \Phi \left( -\sqrt{\alpha} y \right) - t ,
\end{align*}
\]

where $\Phi$ is the cumulative distribution function of a standard normal distribution. By rearranging those expressions, we obtain a necessary and sufficient condition for the “attack” and the “don’t attack” strategy profiles both being equilibria of this game; namely:

\[
y \in \left[ -\frac{\Phi^{-1}(t/D)}{\sqrt{\alpha}}, 1 - \frac{\Phi^{-1}(t/D)}{\sqrt{\alpha}} \right].
\]

Therefore, this incomplete information model may have multiple equilibria or a unique equilibrium depending on whether condition (2) is or is not fulfilled.\(^2\) We can now examine the effects of $y$ and $\alpha$ on both the attack and the no-attack strategy profiles, irrespective of the number of equilibria. These effects are summarized by the following proposition:

**Proposition 2** (i) Both $u(a_i, a_{-i})$ and $u(a_i, d_{-i})$ are decreasing in $y$. (ii) $u(a_i, a_{-i})$ is decreasing (increasing) in $\alpha$ if $y > 1$ ($y < 1$). (iii) $u(a_i, d_{-i})$ is decreasing (increasing) in $\alpha$ if $y > 0$ ($y < 0$).

**Proof.** From equations (1), differentiating $u(a_i, a_{-i})$ and $u(a_i, d_{-i})$ with respect to $y$ yields:

\[
\begin{align*}
    \frac{d}{dy} u(a_i, a_{-i}) &= -D \sqrt{\alpha} \cdot \phi \left[ \sqrt{\alpha} (1 - y) \right] \\
    \frac{d}{dy} u(a_i, d_{-i}) &= -D \sqrt{\alpha} \cdot \phi \left( -\sqrt{\alpha} y \right) .
\end{align*}
\]

where $\phi$ is the probability density function of a standard normal distribution. Thus, both derivatives are always negative.

Differentiating with respect to $\alpha$ we obtain:

\[
\begin{align*}
    \frac{d}{d\alpha} u(a_i, a_{-i}) &= (1 - y) \frac{D}{2\sqrt{\alpha}} \cdot \phi \left[ \sqrt{\alpha} (1 - y) \right] \\
    \frac{d}{d\alpha} u(a_i, d_{-i}) &= -y \frac{D}{2\sqrt{\alpha}} \cdot \phi \left( -\sqrt{\alpha} y \right) .
\end{align*}
\]

\(^2\)Note that, given $D, t,$ and $y$, changes in $\alpha$ (i.e. changes in speculators’ uncertainty about $\theta$) may produce a shift from a model with multiple equilibria to a model with a unique equilibrium. Hence, one can find examples in which modifications in uncertainty trigger a speculative attack, even if the mean of speculators’ expectations $y$ does not change. This feature of currency crisis games is further analyzed in Sbracia and Zaghini (2001).
Therefore, the derivative of \( u(a_i, a_{-i}) \) is negative (positive), provided that \( y > 1 \) \((y < 1)\); the derivative of \( u(a_i, d_{-i}) \) is negative (positive), provided that \( y > 0 \) \((y < 0)\).

An increase in \( y \), by reducing \( u(a_i, a_{-i}) \) and \( u(a_i, d_{-i}) \), shrinks the range of parameter values for which the attack strategy profile is an equilibrium and enlarges the range of parameter values for which the don’t-attack strategy profile is an equilibrium. In other words, an improvement in the expected fundamental always makes it less likely that the attack strategy profile will be an equilibrium and more likely that the no-attack strategy profile will be.

Proposition 2 also states that the effect of the precision of the public signal, \( \alpha \), depends on the ex-ante expected fundamental \( y \). Specifically, if \( \alpha \) increases and \( y \) is sufficiently good (bad), it becomes less (more) likely that the attack strategy profile will be an equilibrium and more (less) likely that the don’t-attack strategy profile will be. In order to understand this dependence of the effect of \( \alpha \) on \( y \), recall that an increase in \( \alpha \) makes speculators more confident that the fundamental \( \theta \) is in a neighborhood of \( y \). Therefore, when \( y \) is sufficiently good, the increase in \( \alpha \) makes all speculators more confident that the peg will hold, dampening their willingness to attack. Conversely, when \( y \) is sufficiently bad, more precise public signals strengthen speculators’ confidence that the currency will depreciate, driving them to attack the peg.\(^3\)

Thus, despite the differences in the number of equilibria and in the information structure, these results confirm that in the presence of multiple equilibria the ex-ante expected fundamental and the precision of public information have effects comparable to those they have in the unique-equilibrium model of Section 2.2.

\(^3\)Note also that for intermediate values of \( y \) \((0 < y < 1)\), if \( \alpha \) increases, there is a widening of the range of parameters in which both the attack and the don’t attack strategy profiles are equilibria of the game.