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Currency crises and uncertainty about fundamentals

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CURRENCY CRISES AND UNCERTAINTY ABOUT FUNDAMENTALS

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Abstract

This paper extends some theoretical results of Morris and Shin (1998) concerning the role of uncertainty about fundamentals in currency crises and tests their empirical relevance using a novel approach based on the distribution of survey expectations. Econometric evidence from the Asian crisis confirms the prediction that the dispersion of expectations affects the probability of a speculative attack and that the sign of this effect depends on whether expected fundamentals are "good" or "bad." Extensive robustness checks support the findings.

JEL classification: F31, D84, D82.

Keywords: speculative attack, exchange rate crisis, public and private information.

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1. Introduction[1]

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 in currency crises 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.[2] 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).

Whereas almost no empirical paper on currency crises has made uncertainty about fundamentals its central focus, some have developed theoretical models in which the variance of fundamentals plays a role. In stochastic "first-generation" models of currency crises, for example, the variance of fundamentals affects the probability of a speculative attack at each point in time (Flood and Garber, 1984). In this class of models, greater uncertainty about fundamentals tends to increase the probability of a speculative attack as long as certain conditions are satisfied.[3] In a recent paper, Flood and Marion (2000) extend Flood and Garber

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

[3] In Goldberg (1991), domestic credit growth follows a random walk process with errors distributed as a

(1984) 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.[4]

"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 matters (see, for example, Obstfeld (1996)). In a second-generation model of currency crises with incomplete information, Sbracia and Zaghini (2001) show that an increase in the variance of public information about fundamentals 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.

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.

Models à la Morris and Shin provide a natural framework for studying the role of uncertainty in currency crises, as private information generates empirically plausible equilibria in which only a fraction of speculators attacks the currency, with or without success. In models with complete – or incomplete but only public – information, only equilibria in mixed strategies could have similar features. In addition, the presence of a unique equilibrium allows

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displaced exponential with zero mean; if the variance of the errors is below an upper bound, greater uncertainty increases the probability of an attack. In counterfactual simulations Goldberg (1994) finds, however, that a higher variance of domestic credit growth would have *reduced* the probability of an attack in Mexico between 1980 and 1986. In Grilli (1986), fundamentals follow an AR(1) process with normal errors; as long as fundamentals are "good," the effect of the variance on the probability of an attack is positive, but with sufficiently "bad" fundamentals it may become negative.

[4] 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.

to perform rigorous comparative statics exercises that are not possible in multiple-equilibrium models à la Obstfeld (1994 and 1996).

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 model with 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 (2002a) question, however, the robustness of this result and, in a somewhat different framework, find that greater precision of information does not always attenuate speculators' coordination problem. Finally, Metz (2002) shows that the effect of changes in the variance of information on the decision rule of the government depends on the expected level of fundamentals and on whether it is the public or the private information precision that varies.[5]

In the first part of the paper, we extend Metz's result in order to obtain predictions about the effects of the precision of public and private information 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. In line with Metz (2002), we find that the effect of public information depends on the expected fundamental: when this is sufficiently good (bad), an increase in the precision of public information makes speculative attacks less (more) likely. In addition, we show that the precision of private information has an effect on the share of attacking speculators similar to that of public information, provided that actual and expected fundamentals are both sufficiently good or both sufficiently bad. Private information is predicted to have a different effect only when actual and expected fundamentals are at odds. To show that our predictions on the effects of the mean and variance of public information would hold *also in the presence of multiple equilibria,* Appendix I considers the special case in which there is no private information.

In the second part of the paper, we verify empirically whether uncertainty about fundamentals contributes to currency crises and whether this effect depends on the level of expected fundamentals, as the theory predicts. The previous empirical literature on exchange rate dynamics has not focused on uncertainty with two main exceptions: Hodrik (1989), who

[5] In a related framework, Corsetti et al. (2002) 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.

has unsuccessfully tried to use 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 have estimated a GARCH-in-Mean restricted VAR model. The mainstream empirical literature on currency crises, including Eichengreen et al. (1996) and Kaminsky and Reinhart (1999), has also generally neglected the role of uncertainty about fundamentals. The focus on the role of uncertainty in currency crises then distinguishes our paper from this previous empirical literature. Moreover, in order to explain exchange rate pressures, we use forward-looking survey forecasts of fundamentals from Consensus Economics rather than only the current level of fundamentals. This paper also provides the first empirical test of models of currency crises à la Morris and Shin. Our results confirm the theoretical predictions that both the mean and the variance of GDP forecasts contribute to explaining exchange rate pressures and that the effect of the variance depends on the level of expected fundamentals.

The paper is organized as follows. After introducing a benchmark model with complete information and multiple equilibria, Section 2 presents the main theoretical results for the model with incomplete public and private information and a unique equilibrium. Section 3 derives the testable implications of the latter, relating its predictions to Consensus Economics forecasts of fundamentals. Section 4 presents the results of our estimates of exchange rate pressures in six Asian countries (Thailand, South Korea, Indonesia, Malaysia, Singapore, and Hong Kong) for the period January 1995 - May 2001, for which Consensus Economics forecast data are available. Section 5 concludes.

2. Theoretical background

Our simple formulation of a currency crisis game builds on Morris and Shin (2002a) and Metz (2002). We first consider a complete information model with multiple equilibria. We then assume that speculators receive both public and private information and characterize the unique equilibrium that emerges when private information is sufficiently precise.

Our incomplete information analysis focuses on the effects of changes in three key parameters: the mean of speculators' expectations about the fundamentals and the precisions of public and private information.[6] An improvement in the *mean* of speculators' expectations

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[6] Appendix I shows that changes in the first two parameters tend to have comparable effects in a model

always makes speculative attacks less likely. The effect of the *precision of public information* depends, instead, on the expected fundamental: when this is sufficiently good (bad), an increase in the precision of public information makes speculative attacks less (more) likely. By extending previous results, we find that the *precision of private information* has two distinct effects. First, it affects the likelihood of an attack *directly*, since it is inversely related to the dispersion of speculators' private signals around the actual fundamental. Second, it affects the likelihood of an attack *indirectly*, as the ratio between the precision of public and private information represents the extent to which speculators expect their beliefs to be shared by other speculators, thereby influencing their "degree of aggressiveness." We show that while these two effects have opposite consequences on the likelihood of an attack, the net effect of private information tends to be similar to that of public information, provided that actual and expected fundamentals are either both sufficiently good or both sufficiently bad; otherwise, the effect of private information precision turns out to be opposite to that of public information precision.

2.1 *Complete information model*

Let us consider a continuum of speculators in an economy characterized by a state of fundamentals θ that can take values over the real line $\Re$, 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.[7] If speculators refrain from attacking, they get 0. The government's utility from defending the currency is increasing in the fundamental θ, and decreasing in the share of speculators that attack the currency, denoted by l. Specifically, we assume that the government gets $\theta - l$, when he maintains the peg and zero when he abandons it.

We consider a very simple two-stage game with complete information. In the first stage, speculators observe θ and simultaneously decide whether to attack the currency. In the second

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where there is only public information and multiple equilibria are possible.

[7] Here we take D constant. Assuming that D depends on the level of fundamentals θ (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.

stage, the government – who knows θ – 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:[8]

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

Given ψ, 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 θ lies in:[9]

- $(-\infty, 0] \Longrightarrow$ there is a unique equilibrium: all agents attack the currency and the government devalues;
- $(0, 1] \Longrightarrow$ 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);
- $(1, +\infty) \Longrightarrow$ 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 θ. By contrast, when θ 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*

We now assume that speculators do not know the fundamental θ but only have expectations about it, given by the following normal probability distribution

$$\Theta \sim Norm(y, 1/\alpha) \,,$$

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[8] We assume – without altering the analysis – that the government chooses to abandon the peg when he is indifferent.

[9] Hereafter we restrict our attention to pure strategies.

with $\alpha > 0$. Since Θ is common knowledge to all speculators, this probability distribution represents the *public information* available to them. Thus, we will refer to α as the "precision of public information." Suppose also that each speculator i receives a *private signal* x_i drawn from the following normal distribution

$$X_i \mid \theta \sim Norm(\theta, 1/\beta) ,$$

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

In this paper, we do not use the term *public information* 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. In this framework, an increase in the variance of the distribution of public information does *not* necessarily reflect noisier official information but it could be due to greater uncertainty – common to all agents – about the economic outlook.[10] 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 that 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 this paper, we do not distinguish between these two sources of uncertainty, both of which would affect the precision of public information. An implication of this approach is that, unlike Morris and Shin (2002b), we do not perform welfare analysis on the provision of public information.

Private information in economic models may arise from a variety of sources. In general, a noisy private signal may represent discrepancies in how public information is interpreted

[10] The sharp increase in the dispersion of GDP forecasts in the aftermath of currency crises documented in Figure 2 may 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.

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 also produce heterogeneity in speculators' information sets, as documented by Kaufmann et al. (2000). Moreover, on foreign exchange markets, international banks may gather valuable private information from monitoring the activity of their customers.

When private information is sufficiently precise with respect to public information, this model entails a unique equilibrium.[11] 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 II illustrates why the lack of common knowledge leads to a unique equilibrium. A condition that grants the existence of a unique equilibrium is:

(1)
$$\beta > \frac{\alpha^2}{2\pi} \, .$$

The intuition for this condition is straightforward. If private signals were not sufficiently informative with respect to the public signal, speculators would regard them as unreliable and continue to ground their decisions mostly on public information, restoring a high degree of common knowledge. Under condition (1), the following proposition holds (see Appendix II):

It is easy to verify that $\theta^* \in [0, 1]$. As a consequence of the unique equilibrium result, the maintenance of the currency peg depends solely on the actual fundamental θ and the parameters y, α, and β. Therefore, speculators' expectations matter, as changes in the mean and in the precisions of public and private information determine the equilibrium trigger points θ^* and x^*. 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]

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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 result in a unique equilibrium. In other words, for the unique-equilibrium result it is also crucial that there be a non-negative probability of θ belonging to $(-\infty, 0)$ and to $(1, +\infty)$. This condition is fulfilled when we assume normal distributions.

[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, which is a stronger equilibrium concept. For instance, in a related framework Heinemann and Illing (2002) exploit this property to show that the

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

2.2.1 *Expectation effects on the equilibrium*

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

The three conditions for α to reduce θ^* and x^*, for β to raise θ^*, and for β to raise x^* respectively are:[13]

$$y > \theta^* - \frac{1}{2\sqrt{\alpha+\beta}}\Phi^{-1}\left(\frac{t}{D}\right) \equiv s_1 \tag{2}$$

$$y > \theta^* - \frac{1}{\sqrt{\alpha+\beta}}\Phi^{-1}\left(\frac{t}{D}\right) \equiv s_2 \tag{3}$$

$$y > \theta^* - \frac{\alpha^2\phi - 2\sqrt{\beta}\alpha - \left(\sqrt{\beta}\right)^3}{\alpha\sqrt{(\alpha+\beta)}\left(\alpha\phi - \beta\phi - 2\sqrt{\beta}\right)}\Phi^{-1}\left(\frac{t}{D}\right) \equiv s_3\ . \tag{4}$$

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

Proposition 1 (Metz, 2002) *Assume that $\beta > \frac{\alpha^2}{2\pi}$. Then θ^* is: (i) decreasing in y; (ii) decreasing (increasing) in α if $y > s_1$ ($y < s_1$); (iii) increasing (decreasing) in β if $y > s_2$ ($y < s_2$).*

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introduction of sunspots (correlation devices unrelated to fundamentals) does not restore multiplicity of equilibria.

[13] Note that, if $D = 2t$, the conditions (2), (3), and (4) coincide.

The effects of the parameters on the decision rule of speculators (i.e. on the trigger point x^*), which are crucial to the theoretical results on the share of attackers presented in the next section, are given by the following proposition (see Appendix III):

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

The effects of the parameters on θ^* and x^* are essentially the same. The most striking result from these propositions concerns the opposite effects of α and β. Key to this result 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 α and β, 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, in fact, his expected fundamental is

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

Suppose that y is sufficiently high (i.e., conditions (2)-(4) hold) so that speculators will on average expect "good" fundamentals. In this situation, if the precision ratio $\frac{\alpha}{\beta}$ is also high, speculators know that other speculators have formed their 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"; in this case, relatively precise private (public) information helps (hurts) the government by making it harder (easier) for speculators to coordinate the attack on the currency.[14]

2.2.2 *Expectation effects on the share of attackers*

We can now derive the effects of the parameters on the share of speculators attacking the currency, which is equal to the probability of one speculator attacking, $\Pr(X_i \leq x^* \mid \theta)$, or $\Phi\left[\sqrt{\beta}(x^* - \theta)\right]$. This probability depends on the actual fundamental θ and on the parameters y, α and β. Given the results in Proposition 3, by differentiating $\Phi\left[\sqrt{\beta}(x^* - \theta)\right]$ it is easy to obtain:

Proposition 3 *Assume that* $\beta > \frac{\alpha^2}{2\pi}$*; then the probability* $\Pr(X_i \leq x^* \mid \theta)$ *is: (i) decreasing in* θ*; (ii) decreasing in* y*; (iii) decreasing (increasing) in* α *if* $y > s_1$ *(*$y < s_1$*). (iv) decreasing (increasing) in* β *if* $\frac{(x^*-\theta)}{2\sqrt{\beta}} + \sqrt{\beta}\frac{dx^*}{d\beta} < 0$ *(*$\frac{(x^*-\theta)}{2\sqrt{\beta}} + \sqrt{\beta}\frac{dx^*}{d\beta} > 0$*).*

As expected, an improvement in θ or in y reduces the share of speculators attacking the currency. Similarly, the effect of the precision of public information is in line with previous results, since it only depends on the expected fundamental y: when y is sufficiently good (bad), an increase in α causes the share of speculators to decrease (increase). However, unlike the prediction of Propositions 2 and 3, the effect of the precision of private information is not necessarily opposite to that of public information. In order to understand this new result, we need to consider that β not only affects the equilibrium trigger point x^* (thus indirectly influencing the share of speculators attacking the currency) but also directly affects the probability of receiving a private message smaller than x^*, $\Pr(X_i \leq x^* \mid \theta)$, since it determines the dispersion of the messages x_i around the actual fundamental θ.

Consider the following example. Assume that y is good ($y > s_3$), so that an increase in β causes x^* to increase ($\frac{dx^*}{d\beta} > 0$), making speculators more aggressive for any θ. One might

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[14] Heinemann and Illing (2002) obtain a different result on the effect of private information: in their model an increase in the precision of private information, β, always decreases θ^*, making speculative attacks less likely. However, Heinemann and Illing assume that θ 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 β always reduces θ^*, provided that condition (3) 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 θ. This means that if the attack is successful and θ 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 θ more likely – tends to drive speculators to the attack strategy.

expect that the share of speculators attacking the currency would also increase. Nevertheless, if the actual fundamental θ is sufficiently good ($\theta > x^*$), the increase in β reduces the dispersion of the distribution around the good fundamental, so a larger number of speculators receive good signals. When this second effect is strong enough, it offsets the indirect effect of β on x^*, and the resulting share of attackers decreases. Figure 3 illustrates of the effect of an increase in β for a good value of y. When β rises, the resulting increase in x^* (indirect effect) tends to increase the share of attackers for each value of θ (dotted line). However, as a consequence of the direct effect, the slope of the curve changes, so that the share of attackers increases only for low values of θ and decreases for high ones.

In general, the net effect of an increase in β on the share of speculators depends on the relative strength of these two effects. When the direct effect prevails, the effect of the precision of private information is analogous to that of public information. Note, also, that the direct effect tends to prevail either when θ and y are both sufficiently good or when they are both sufficiently bad. Conversely, the effect of the precision of private information tends to be opposite to that of public information when the indirect effect dominates, which occurs when θ is good and y is bad or vice versa.

3. Testable implications

We use forecasts of fundamentals collected by Consensus Economics to verify whether the mean and variance of agents' expectations contribute to explaining actual exchange rate pressures. 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 of Section 2.2, we assume that each forecaster declares to Consensus Economics the mean of his posterior probability distribution. (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.[15]) Recall that, given the message x_i, the posterior probability distribution is

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

[15] A formal proof is available from the authors upon request.

Our assumption implies that the prediction about the fundamental θ that agent i (i.e., the agent receiving the message x_i) reports to Consensus Economics is the posterior mean $f_i^e(x_i)$ given by equation (5). Let us consider the mean of the individual forecasts, i.e.

$$f^e(x_1, ..., x_n) = \frac{\sum f_i^e(x_i)}{n} = \frac{\alpha}{\alpha+\beta}y + \frac{\beta}{\alpha+\beta}\frac{\sum x_i}{n} \tag{6}$$

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

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

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 θ and y. Recall that θ and y have the same effects on the share of attackers: when actual or expected fundamentals improve (deteriorate), pressures on the exchange rate will abate (strengthen). Note also that the theoretical model suggests that $E[f(\Theta)] = y$; thus, *on average* the mean of individual forecasts is equal to the expected fundamental y and does not depend in any systematic way on α and β. Similarly, in our empirical work we expect that, *along the time-series dimension*, the mean of individual forecasts does not depend on α and β.

The theoretical model also implies that the precisions of public and private information affect exchange rate pressures (points *(iii)-(iv)* in Proposition 4) in a way that is distinct from that of actual and expected fundamentals (points *(i)-(ii)* in Proposition 4). In other words, Proposition 4 suggests that even if actual and expected fundamentals remain unchanged, speculative pressures on the exchange rate vary with the variance of public or private information. Empirically, changes in the precision of public and private information will be reflected in the variance of the individual forecasts:

$$[\sigma^e(x_1, ..., x_n)]^2 = \sum \frac{[f_i^e(x_i) - f^e(x_1, ..., x_n)]^2}{n} = \frac{\beta^2}{(\alpha+\beta)^2}\frac{\sum(x_i - \overline{x})^2}{n} \, , \tag{7}$$

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

$$\sigma^2 = E\left[(\sigma^e(X_1, ..., X_n))^2 \mid \theta\right] = \frac{\beta}{(\alpha+\beta)^2} \, . \tag{8}$$

Hence, for n sufficiently large, a change in y affects the individual forecasts f_i^e but does not affect their variance σ^2, which only depends on α and β. 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 (8) that while an increase in α always implies a decrease in σ^2, an increase in β does not necessarily reduce σ^2. This result is easily explained. On the one hand, β tends to reduce σ^2 as it decreases the dispersion of the messages x_i, but on the other hand, for given messages x_i, the rise in β increases the weight of the private messages in the individual predictions (5), 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 σ^2 is decreasing in β, so that we can always interpret a decline in σ^2 as due to an increase in either α or β or both. The condition $\beta > \frac{\alpha^2}{2\pi}$ ensures the existence of a unique equilibrium and that Proposition 4 holds.[16]

From Proposition 4 we know that the effect of σ^2 on speculative pressures will depend on whether it is α or β that changes and on the level of the expected fundamental y. We therefore estimate a specification of the following general form:

$$ERP_t = \gamma_0 + \gamma_1 f_{t-1}^e + \gamma_2 \sigma_{t-1}^e \cdot (f_{t-1}^e - \underline{\gamma}) + \gamma_3 e_{t-1} + \varepsilon_t \tag{9}$$

where ERP is a measure of exchange rate pressure, f^e is the mean of the individual forecasts from (6), σ^e is the standard deviation corresponding to the square root of (7), $\underline{\gamma}$ is the threshold separating "good" from "bad" expected fundamentals, and e is the real effective exchange rate. All regressors are lagged one period to avoid simultaneity bias.

We expect the coefficient γ_1 to be negative because an improvement in the expected level of fundamentals eases the pressure on the exchange rate.

The effect of an increase in the dispersion of the individual forecasts, σ^e, depends on the expected fundamental and on the source of uncertainty. The parameter $\underline{\gamma}$ is the empirical

[16] Appendix I 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.

proxy for the right-hand side of equations (2) and (4). If changes in the precision of *public* information are at the origin of most changes of σ^e in our sample, then γ_2 should be positive, because by Proposition 4 imprecise public information (i.e., a high σ^e due to a low α) with good expected fundamentals (i.e., $f^e_{t-1} > \underline{\gamma}$) heightens exchange rate pressures. We also expect γ_2 to be positive if the changes in σ^e are due to changes in the precision of *private* information while actual and expected fundamentals are either both sufficiently good or both sufficiently bad, so that the *direct effect* of β on the likelihood of an attack dominates (see previous section). In principle, γ_2 could be negative if changes in σ^e were due to changes in the precision of *private* information and actual and expected fundamentals were sufficiently different in most of the sample (i.e., either the actual fundamental is good and the expected fundamental is bad or the actual fundamental is bad and the expected fundamental is good).

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 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 γ_3 to be positive.

4. Empirical evidence

In this section, we verify whether the mean and variance of agents' expectations for economic fundamentals help to 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 verify whether expected fundamentals and their dispersion affect the fraction of speculators that decide to attack the currency, we build an index of exchange rate pressure.[17]

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[17] Girton and Roper (1977), Roper and Turnovsky (1980), and Weymark (1998) discuss the assumptions

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 $IND3$ 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, we also compute an index $IND2$, which sums only normalized values of *i)* and *ii)*,[20] and an index BIS, which is the continuous version of an index recently developed by the Bank for International Settlements for monitoring purposes.[21] Figure 4 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. Following Brooks et al. (2001), in order to reproduce a

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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.

[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.

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.[22] 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 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 (9):

$$\begin{aligned} IND3_{j,t} &= \hat{\gamma}_{0,j} + \hat{\gamma}_1 f^e_{GDP_j,t-1} + \hat{\gamma}_2 \sigma^e_{GDP_j,t-1} \cdot (f^e_{GDP_j,t-1} - \underline{\hat{\gamma}_j}_{GDP}) \\ &+ \hat{\gamma}_{3,j} e_{j,t-1} + u_{j,t} \qquad , \qquad u_{j,t} = \hat{\rho}_j u_{j,t-1} + \varepsilon_{j,t} \end{aligned} \tag{10}$$

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

[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, the current-year forecast, or the following-year forecast together with the difference between the two. In these cases, the dispersion measures have been 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.

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 (10). 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 $\hat{\gamma}_{1,j}$ ($j = 1, .., 6$) were negative and statistically significant at the 5 percent confidence level and all $\hat{\gamma}_{2,j}$ ($j = 1, .., 6$) were positive and statistically significant at the 1 percent confidence level. Note also that uncertainty about GDP growth contributes considerably to explaining exchange rate pressures: if we set $\hat{\gamma}_2 = 0$ in equation (10), the R^2 for the overall system falls from 42.6 to 33 percent.

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 results are consistent with the hypothesis that changes in the precision of public information are the main factors behind the time-series variation in the dispersion of the forecasts; or, if it is the precision of private signals that varies, that the direct effect of precision changes on the distribution of the signals dominates the indirect effect on the trigger point x^* (see Proposition 4).

4.3 *Sensitivity analysis*

This section presents a series of robustness tests of our benchmark specification (10), confirming our main result that uncertainty about fundamentals plays a role in currency crises and that this role depends on the expected level of the fundamentals.

4.3.1 *Robustness to alternative exchange rate pressure measures.*

Tables 2 and 3 present estimates of the specification (10) with two alternative measures of exchange rate pressure as dependent variable ($IND2$ and 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.

4.3.2 *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.[23] 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 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.

[23] However, during the crisis Singapore claimed to have broadened the undisclosed target band within which the Singapore dollar was allowed to fluctuate.

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 $\underline{\hat{\gamma}}$, $\hat{\gamma}_3$, and $\hat{\rho}_j$ to be the same in all countries, allowing only the intercepts $\hat{\gamma}_{0,j}$ 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.[24] 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.

4.3.3 *Robustness to dynamic specification and spurious correlation.*

The model in Section 2.2 is static, then to correct for serial correlation of the errors, rather than estimating a dynamic specification, we included a country-specific AR(1) term in our benchmark regression (10). In a possible dynamic extension of our theoretical model,

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[24] Jeanne and Rose (2002) show, for example, that market expectations should be noisier under a floating exchange rate regime.

however, past values of the exchange rate pressure index could contain information about the stochastic process generating the fundamentals, which speculators would then include in their learning processes. The empirical counterpart of this theoretical model would be a dynamic regression specification with the lagged exchange rate pressure index on the right-hand side. Estimates of a dynamic version of equation (10) yielded results very similar to those reported in Table 1 confirming sign and statistical significance of all coefficients.

More generally, correcting for serial correlation or including a lagged dependent variable rules out the possibility that our results might be driven by spurious correlation between the exchange rate pressure index and uncertainty about fundamentals.[25] The spurious regression problem would emerge if the exchange rate pressure were serially correlated and the uncertainty were a function of exchange rate pressures. In this case, the estimated coefficient on the lagged variance of GDP forecasts would mainly reflect the serial correlation of the exchange rate pressure series. As shown in Hamilton (1994, pp. 561-562), correcting for serial correlation or including a lagged dependent variable overcomes the potential spurious regression problem.

4.3.4 *Time-varying $\hat{\gamma}_1$ and $\hat{\gamma}_2$.*

Another robustness check regards the coefficients $\hat{\gamma}_1$ and $\hat{\gamma}_2$. Proposition 4 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 5 (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 θ 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 5 (bottom panel) shows that

[25] This problem is distinct from the possible simultaneous feedback effect of exchange rate pressures onto the mean and variance of fundamentals, which would cause a potential endogeneity problem that we address by lagging all regressors.

the recursive estimated $\hat{\gamma}_2$ changes over time but remains always positive and significantly different from zero.[26]

4.3.5 *Time-varying threshold* $\underline{\hat{\gamma}}$.

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 $\underline{\hat{\gamma}_j}$ in (10) recursively (Figure 6). 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 $\sigma^e_{GDP_j,t-1}$ on exchange rate pressures (measured by $\hat{\gamma}_2 \cdot (f^e_{GDP_j,t-1} - \underline{\hat{\gamma}_j}_{GDP,t-1}))$ may also vary with changes in GDP forecasts ($f^e_{GDP_j,t-1}$) and country-specific thresholds ($\underline{\hat{\gamma}_j}_{GDP,t-1}$). Figure 7 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. Conclusion

This paper studies how uncertainty about fundamentals contributes to currency crises, both theoretically and empirically. The theoretical model shows that speculative attacks depend not only on the current and the expected level of fundamentals but also on the variance of speculators' expectations about fundamentals. This variance affects exchange rate pressures in different ways depending on the level of current and expected fundamentals and on whether it is public or private information that varies in degree of precision. Specifically, if the expected fundamental is sufficiently good (bad), then an increase in the precision of *public* information makes speculative attacks less (more) likely. The effect of the precision of *private* information is twofold: it affects the likelihood of an attack *directly*, since it is inversely related to the dispersion of speculators' private signals around the actual fundamental, and also *indirectly*,

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[26] 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 6 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.

as the ratio between the precision of public and private information represents the extent to which speculators expect their beliefs to be shared, thereby influencing their 'aggressiveness'. We find that while these two effects have opposite consequences on the likelihood of an attack, the net effect of the precision of private information tends to be similar to that of public information when actual and expected fundamentals are either both sufficiently good or both sufficiently bad. The precision of private information can have an opposite effect only if actual and expected fundamentals are at odds, which is unlikely to happen on a regular basis.

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 expected rate 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 7). 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 outlook. 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.

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 expected fundamentals are good, but hurts when they are 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.

Appendix I: A model with only public information

In this section we derive the effects of the mean of speculators' expectations and of the precision of public information in a model with only public information. This model is relevant because it implies effects of the mean and precision of public information similar to those of the unique-equilibrium model of Section 2.2, even though multiple equilibria are now possible.[27] Specifically, also in this model, the way in which uncertainty contributes to currency crises depends on the expected level of fundamentals, thereby providing some further theoretical support to the empirical evidence of Section 4.

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

Given ψ, 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:

$$\begin{aligned} u(a_i, a_{-i}) &= \int_{-\infty}^{1} (D-t)\,\eta(\theta)d\theta - \int_{1}^{+\infty} t\eta(\theta)d\theta \\ u(a_i, d_{-i}) &= \int_{-\infty}^{0} (D-t)\,\eta(\theta)d\theta - \int_{0}^{+\infty} t\eta(\theta)d\theta \end{aligned}$$

where η is the probability density function of Θ. The following proposition specifies the Bayesian Nash equilibria of the reduced-form game of speculators:

[27] In 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 is an equilibrium.

Proposition 4 *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 us rewrite the two expected payoffs as:

$$\begin{aligned} u(a_i, a_{-i}) &= D \cdot \Phi\left[\sqrt{\alpha}\,(1-y)\right] - t \\ u(a_i, d_{-i}) &= D \cdot \Phi\left(-\sqrt{\alpha} y\right) - t\,, \end{aligned} \tag{11}$$

where Φ 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:

$$y \in \left[-\frac{\Phi^{-1}(t/D)}{\sqrt{\alpha}}, 1 - \frac{\Phi^{-1}(t/D)}{\sqrt{\alpha}}\right] . \tag{12}$$

Therefore, this incomplete information model may have multiple equilibria or a unique equilibrium depending on whether condition (12) is or is not fulfilled.[28] We can now examine the effects of y and α 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 5 *(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* α *if* $y > 1$ *(*$y < 1$*). (iii)* $u(a_i, d_{-i})$ *is decreasing (increasing) in* α *if* $y > 0$ *(*$y < 0$*).*

Proof. Differentiating $u(a_i, a_{-i})$ and $u(a_i, d_{-i})$ with respect to y, using equations (11) yields:

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

[28] Note that, given D, t, and y, changes in α (i.e. changes in speculators' uncertainty about θ) 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).

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

Differentiating with respect to α we obtain:

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

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 the mean 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 6 also states that the effect of the precision of the public signal, α, depends on the expected fundamental y. Specifically, if α increases and expected fundamentals are 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 α on the mean y, recall that an increase in α makes speculators more confident that the fundamental θ is in a neighborhood of y. Therefore, when y is sufficiently good, the increase in α 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.[29]

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 mean and variance of public information have effects comparable to those they have in the unique-equilibrium model of Section 2.2.

---

[29] Note also that for intermediate values of y $(0 < y < 1)$, if α 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.

Appendix II: 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 2002a) and Metz (2002) have shown that the unique equilibrium can be specified by a couple (x^*, θ^*), such that speculators use a trigger strategy

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

and the government follows the rule:[30]

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

Here, we first assume that agents use a trigger strategy like δ; we then derive a sufficient condition granting that unique values of x^* and θ^* exist; finally, we find the equations that characterize these values. We do not show that under the sufficient condition a trigger strategy for speculators is the unique optimal strategy, as this result follows directly from Morris and Shin (2002a) or, in a more general framework, from Frankel et al. (2002).

[30] Given θ, the share of attackers is completely determined by δ, since we have assumed that there is a continuum of speculators. It follows that when speculators use δ, the function ψ below is exactly the same as the government's decision rule specified in Section 2 (which was therefore denoted by the same symbol ψ).

Cut-off points

Assume that agents use the trigger strategy δ defined above and let us find the trigger point of the government's optimal strategy. Given x^* and θ, the share of speculators attacking the currency is

$$\Pr\left(X < x^* \mid \theta\right) = \Phi\left[\sqrt{\beta}\left(x^* - \theta\right)\right] .$$

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 θ^* that solves:

$$\theta^* - \Phi\left[\sqrt{\beta}\left(x^* - \theta^*\right)\right] = 0 . \tag{13}$$

Equation (13) implicitly defines θ^* as a function of x^*. Note that Φ is decreasing and continuous in θ^*, and takes all the values in the open interval $(0, 1)$. Therefore, there exists a unique value of θ^* that solves (13), for any $x^* \in \Re$.

Let us find the trigger point for speculators. Given ψ, 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:

$$D \cdot \Phi\left[\sqrt{\alpha + \beta}\left(\theta^* - \frac{\alpha}{\alpha + \beta} y - \frac{\beta}{\alpha + \beta} x^*\right)\right] - t = 0 . \tag{14}$$

Sufficient condition for a unique equilibrium

Unlike equation (13), equation (14) does not necessarily have a unique solution. Note that, as x^* goes to $-\infty$, the left-hand side of equation (14) goes to $D - t > 0$; when x^* goes to $+\infty$, the left-hand side of equation (14) goes to $-t < 0$. By the continuity of the left-hand side of (14), a sufficient condition granting that a unique solution to equation (14) exists may be obtained by requiring that the derivative of the left-hand side of (14) with respect to x^* is smaller than zero; namely:

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

The previous inequality holds provided that

$$\frac{d\theta^*}{dx^*} - \frac{\beta}{\alpha+\beta} < 0 \,. \tag{15}$$

Differentiating implicitly equation (13) we can obtain

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

and, substituting into (15),

$$\frac{\sqrt{\beta}}{\frac{1}{\phi\left[\sqrt{\alpha+\beta}\left(\theta^* - \frac{\alpha}{\alpha+\beta} y - \frac{\beta}{\alpha+\beta} x^*\right)\right]} + \sqrt{\beta}} < \frac{\beta}{\alpha+\beta} \,. \tag{16}$$

A sufficient condition for inequality (16) 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 ϕ is $1/\sqrt{2\pi}$ – we obtain the sufficient condition (1).

Equilibrium

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

$$\begin{aligned} 0 &= \theta^* - \Phi \left[\sqrt{\beta} \left(x^* - \theta^* \right) \right] \\ 0 &= D \cdot \Phi \left[\sqrt{\alpha+\beta} \left(\theta^* - \frac{\alpha}{\alpha+\beta} y - \frac{\beta}{\alpha+\beta} x^* \right) \right] - t \,. \end{aligned} \tag{17}$$

Appendix III: Proofs for propositions 2 and 3

In order to derive the effects of the parameters y, α, β on (x^*, θ^*) the system (17) can also be written as

$$
\begin{aligned}
x^* &= \theta^* + \frac{1}{\sqrt{\beta}}\Phi^{-1}(\theta^*) \\
x^* &= \frac{\alpha+\beta}{\beta}\theta^* - \frac{\alpha}{\beta}y - \frac{\sqrt{\alpha+\beta}}{\beta}\Phi^{-1}\left(\frac{t}{D}\right)
\end{aligned} \tag{18}
$$

that, by substitution, yields:

$$
\theta^* = \Phi\left[\frac{\alpha}{\sqrt{\beta}}\left(\theta^* - y - \frac{\sqrt{\alpha+\beta}}{\alpha}\Phi^{-1}\left(\frac{t}{D}\right)\right)\right] . \tag{19}
$$

In the following, by differentiating the system of implicit equations (17) (or the alternative expressions (19)) we simultaneously obtain the effect of each parameter on both θ^* and x^*, thereby proving both propositions 2 and 3.

Effects of y

By differentiating the system of implicit equations (17) with respect to y, we can obtain:

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

where we have neglected the argument of ϕ. Solving by substitution, we get:

$$
\begin{aligned}
\frac{d\theta^*}{dy} &= -\frac{\alpha\phi}{\sqrt{\beta} - \alpha\phi} \\
\frac{dx^*}{dy} &= -\frac{\alpha}{\beta} + \frac{\alpha+\beta}{\beta}\frac{d\theta^*}{dy} .
\end{aligned}
$$

Therefore, the derivative of θ^* 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 α

In order to derive the effect of α on θ^*, we can simplify our calculations starting by differentiating equation (19):

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

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

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

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 θ^* with respect to α is negative, provided that condition (2) holds.

Let us turn to the effect of α on x^*. Differentiating the first equation of system (17) with respect to α 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 θ^*.

Effect of β on θ^*

Let us differentiate equation (19) with respect to β:

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

where we have neglected the argument of ϕ. Solving by substitution, we get:

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

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 θ^* with respect to β is positive, provided that condition (3) holds.

Effect of β on x^*

Consider the second equation in system (18) and differentiate it with respect to β:

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

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

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

Figures and tables

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

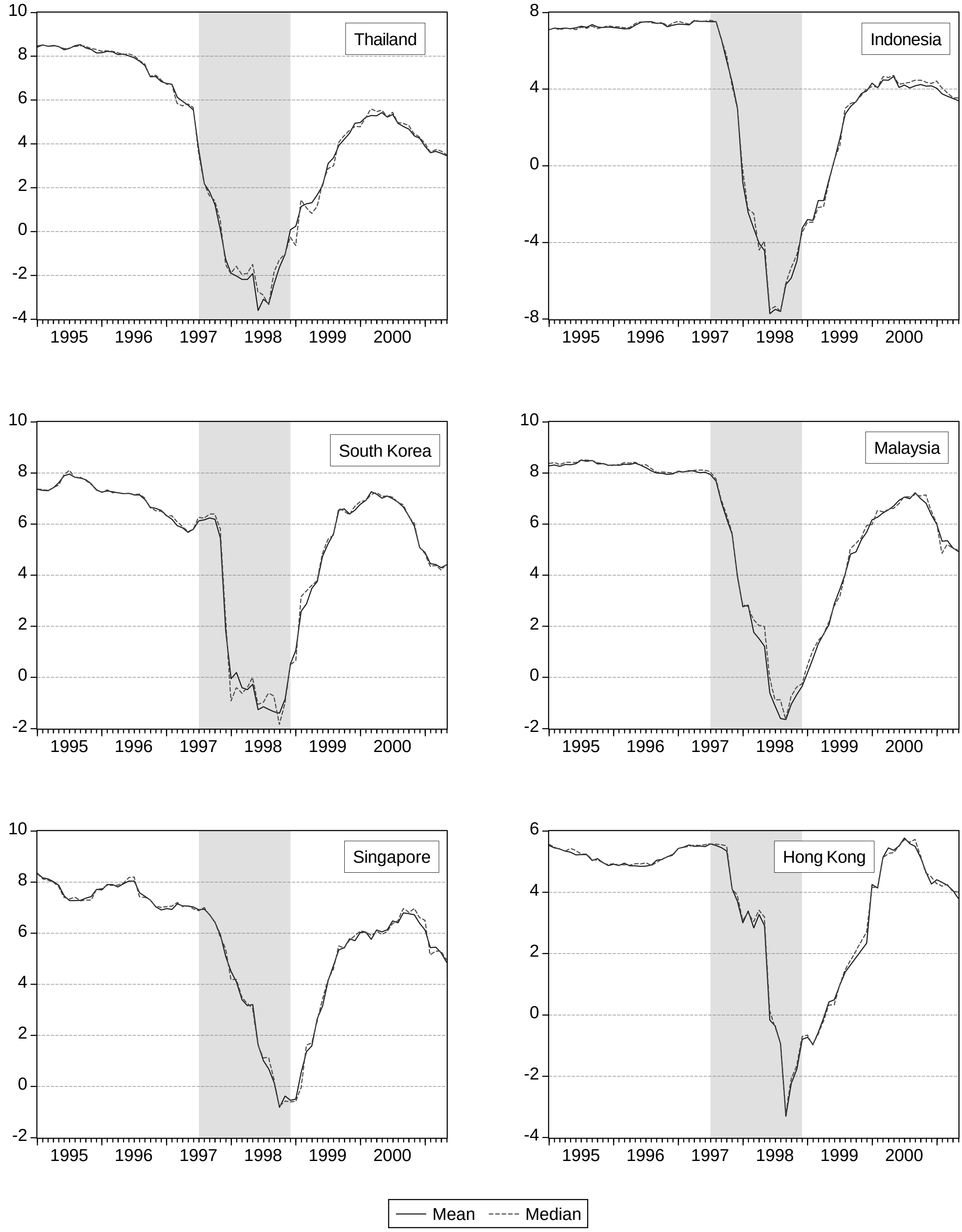


Figure 2: Standard deviation and mean absolute median difference of GDP growth
(weighted average of current and following year forecasts; 1995:01-2001:05)

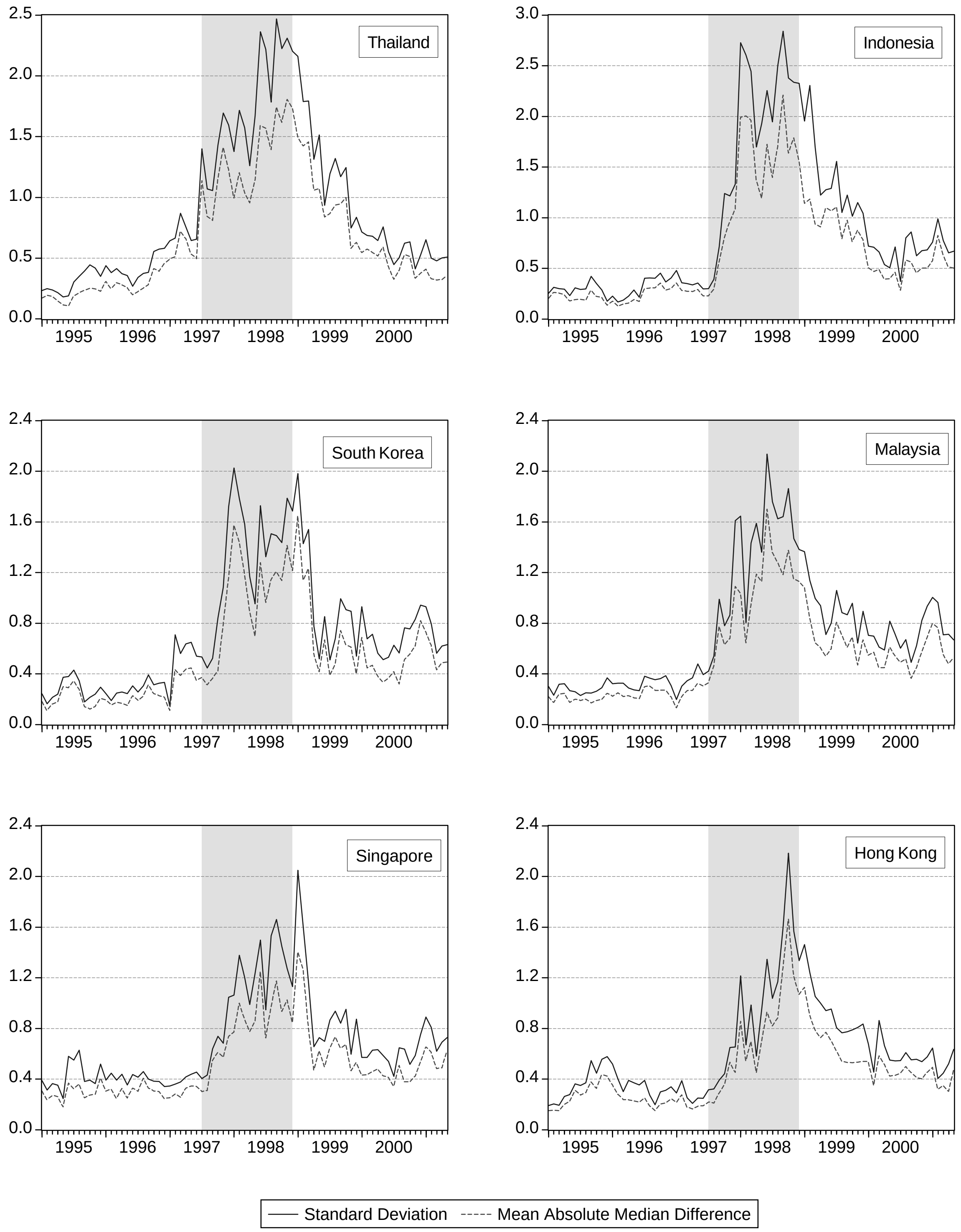
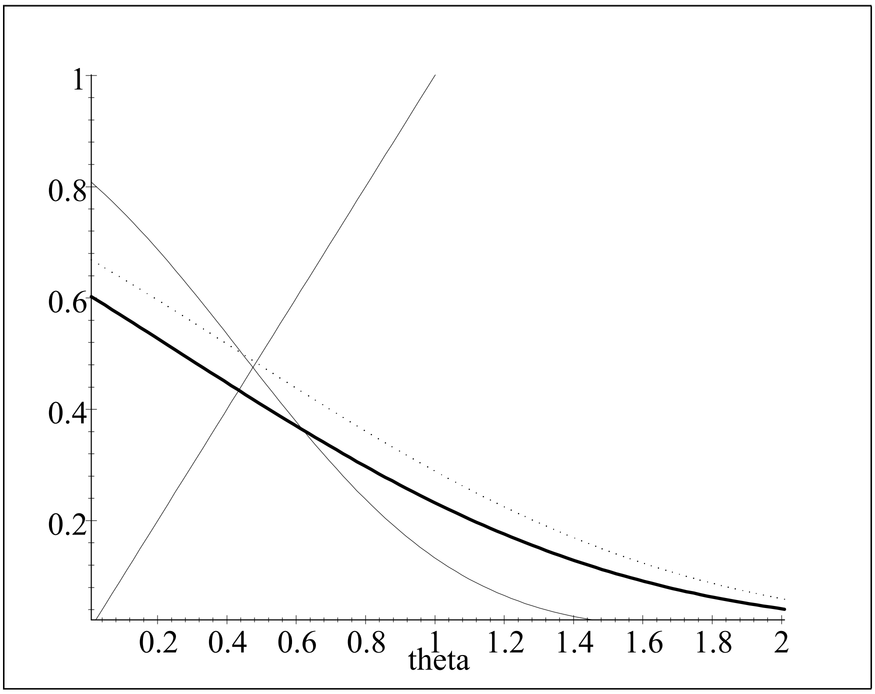


Figure 3: Effects of an increase in β for y "good"



Note: the thick line is the share of attackers (as a function of θ) 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 β increased to 4 (x^* raises to 0.444). The dotted line singles out the *indirect effect* of β as it shows the share of attackers with the new $x^* = 0.444$ and the old $\beta = 1$.

Figure 4: Indices of exchange rate pressure
(1995:01 - 2001:05)

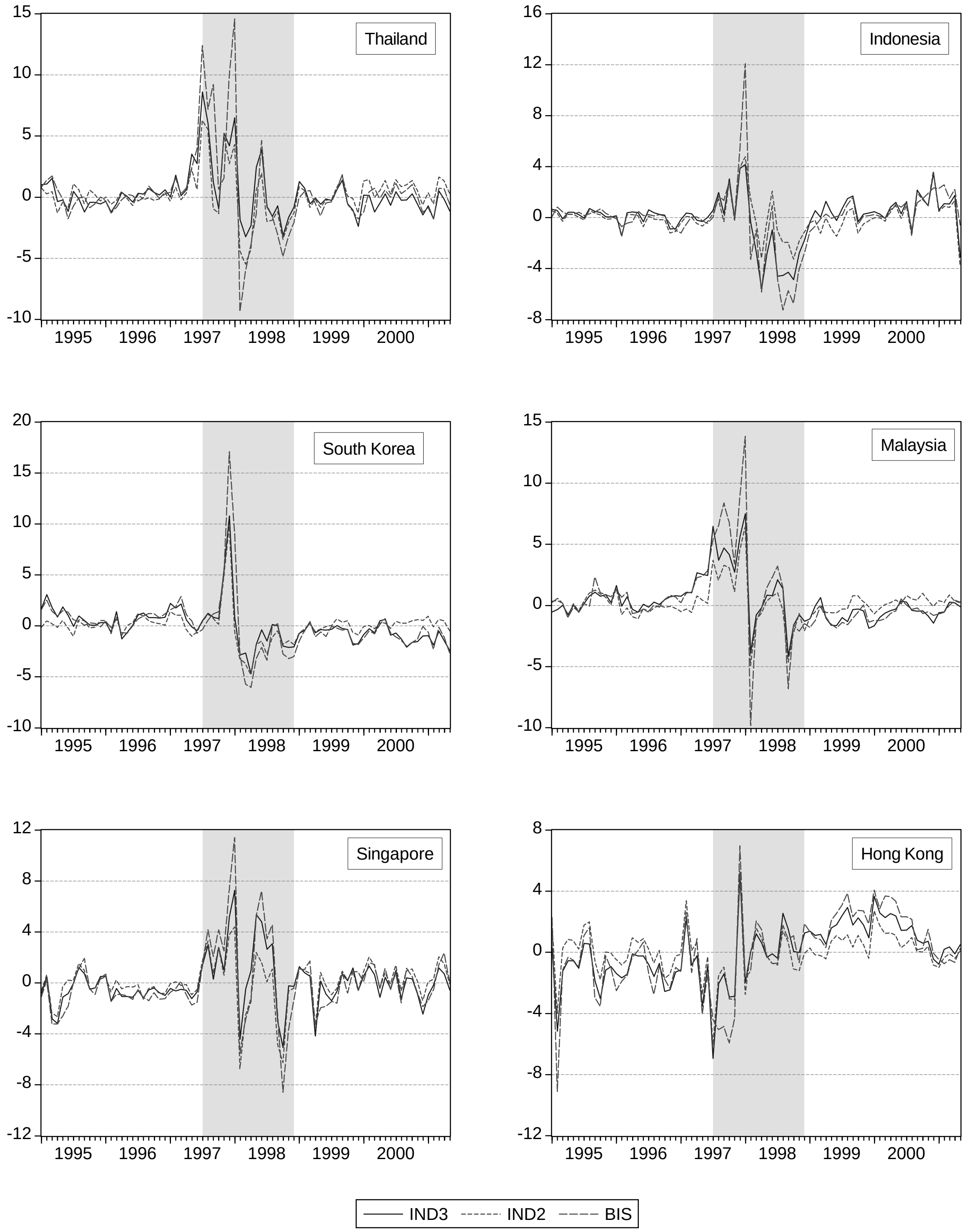
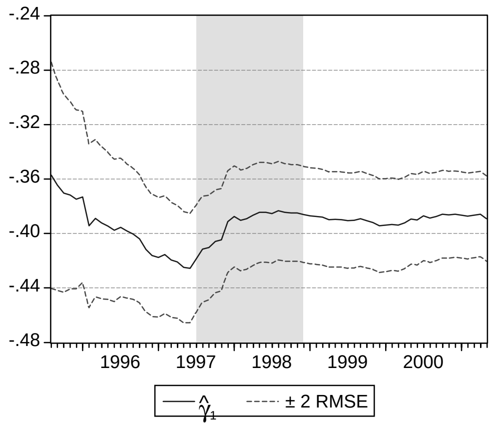


Figure 5: Recursive estimates of $\hat{\gamma}_1$ and $\hat{\gamma}_2$
(1995:08-2001:05)



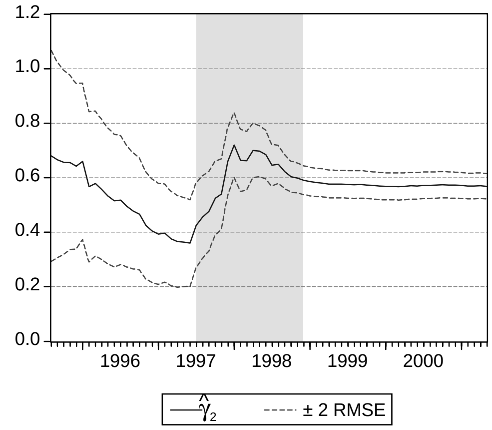


Figure 6: Recursive estimates of the threshold separating high from low expected GDP growth (1996:07 - 2001:05)

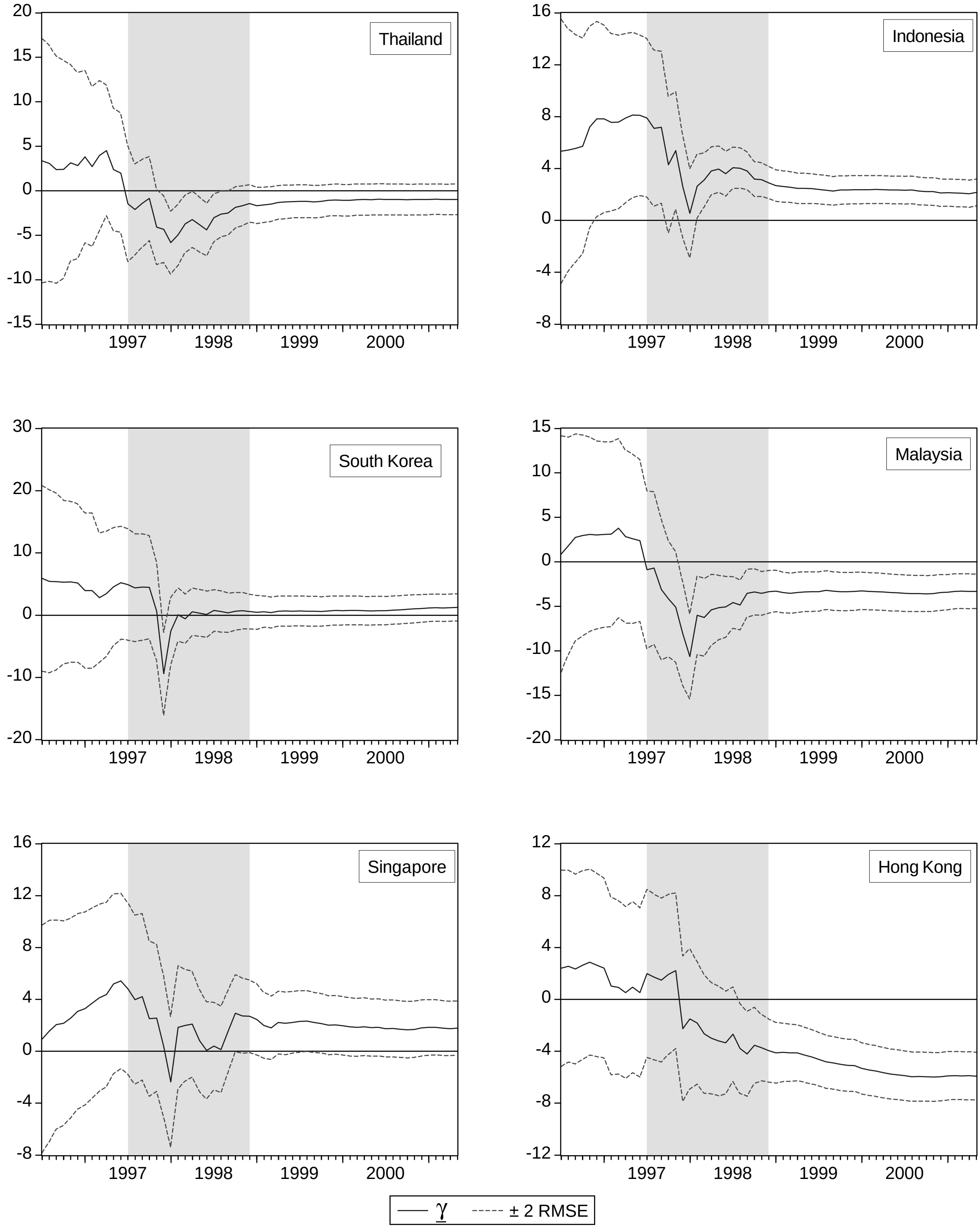


Figure 7: Overall effect of uncertainty on exchange rate pressures in estimates with recursive threshold (1996:07-2001:05)

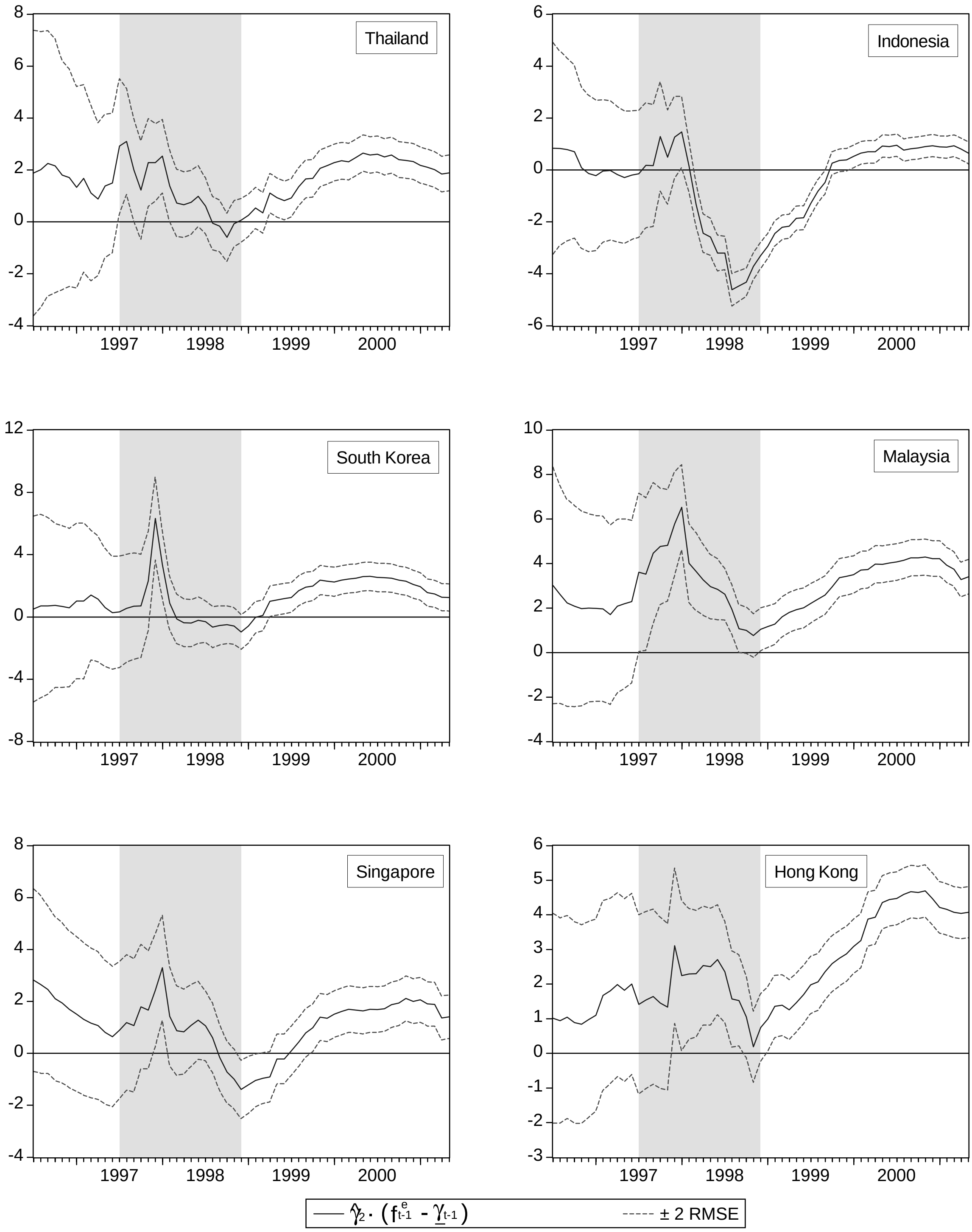


Table 1. Exchange Rate Pressure (**IND3** Index) Estimates
(SUR estimates; standard errors in parenthesis; sample: 1995:03-2001:05)[1]

	Thailand	*Indonesia*	*South Korea*	*Malaysia*	*Singapore*	*Hong Kong*
$\gamma_{0,j}$	-10.645 ***	-1.654	-17.254 ***	-13.830 ***	-18.052 ***	6.885 **
	(3.234)	(1.347)	(3.521)	(2.318)	(6.509)	(3.439)
γ_1	-0.520 ***					
	(0.085)					
γ_2	0.592 ***					
	(0.066)					
γ_j	1.360	1.151	0.495	0.160	3.468 ***	-0.710
	(0.971)	(0.946)	(1.312)	(1.249)	(1.310)	(1.439)
$\gamma_{3,j}$	0.142 ***	0.047 ***	0.238 ***	0.165 ***	0.184 ***	-0.046 *
	(0.035)	(0.015)	(0.042)	(0.022)	(0.058)	(0.027)
ρ_j	0.340 ***	0.283 **	0.503 ***	0.345 ***	0.203 **	0.294 ***
	(0.088)	(0.125)	(0.090)	(0.082)	(0.085)	(0.092)
R^2	0.330	0.647	0.518	0.475	0.227	0.389
DW	1.626	1.971	1.538	1.806	1.695	2.276
Observations	76	76	76	76	76	76

[1] Data are monthly.
Three (***), two (**), and one (*) stars mark statistical significance respectively at one, five, and ten percent levels.
The coefficients γ_1 and γ_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)[1]

	Thailand	*Indonesia*	*South Korea*	*Malaysia*	*Singapore*	*Hong Kong*
$\gamma_{0,j}$	-6.720 ***	-1.193	-11.187 ***	-4.630 ***	-8.223	6.385 **
	(2.569)	(1.528)	(2.865)	(1.775)	(5.484)	(2.719)
γ_1	-0.203 **					
	(0.081)					
γ_2	0.333 ***					
	(0.067)					
γ_j	0.954	-1.582	-0.592	0.933	4.719 **	0.328
	(1.447)	(1.699)	(2.204)	(1.857)	(2.140)	(2.162)
$\gamma_{3,j}$	0.083 ***	0.018	0.144 ***	0.056 ***	0.084 *	-0.048 **
	(0.029)	(0.017)	(0.034)	(0.017)	(0.049)	(0.022)
ρ_j	0.173 **	0.308 **	0.304 ***	0.138	0.089	0.135
	(0.088)	(0.126)	(0.098)	(0.087)	(0.086)	(0.090)
R^2	0.233	0.281	0.440	0.221	0.087	0.121
DW	1.590	1.912	1.500	1.706	1.593	2.252
Observations	76	76	76	76	76	76

[1] Data are monthly.
Three (***), two (**), and one (*) stars mark statistical significance respectively at one, five, and ten percent levels.
The coefficients γ_1 and γ_2 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]

	Thailand	*Indonesia*	*South Korea*	*Malaysia*	*Singapore*	*Hong Kong*
$\gamma_{0,j}$	-30.862 ***	-7.491 ***	-26.437 ***	-23.790 ***	-26.928 ***	19.706 ***
	(7.044)	(2.184)	(4.356)	(3.528)	(7.693)	(7.097)
γ_1	-0.641 ***					
	(0.130)					
γ_2	0.623 ***					
	(0.094)					
γ_j	1.511	-0.702	0.505	-1.660	5.530 ***	2.820
	(1.254)	(1.080)	(1.076)	(1.746)	(1.532)	(2.018)
$\gamma_{3,j}$	0.374 ***	0.117 ***	0.361 ***	0.269 ***	0.275 ***	-0.136 **
	(0.076)	(0.026)	(0.054)	(0.034)	(0.069)	(0.056)
$\rho_{1,j}$	0.568 ***	0.443 ***	1.076 ***	0.330 ***	0.514 ***	0.510 ***
	(0.096)	(0.102)	(0.083)	(0.081)	(0.075)	(0.088)
$\rho_{2,j}$	-0.123 *	0.148 *	-0.452 ***	0.037	-0.255 ***	0.170 **
	(0.072)	(0.081)	(0.069)	(0.074)	(0.071)	(0.078)
R^2	0.258	0.452	0.599	0.323	0.297	0.353
DW	1.815	2.170	1.853	1.894	1.959	2.245
Observations	76	76	76	76	76	76

[1] Data are monthly.
Three (***), two (**), and one (*) stars mark statistical significance respectively at one, five, and ten percent levels.
The coefficients γ_1 and γ_2 are restricted to be the same across countries.

Table 4. Exchange Rate Pressure Estimates on Pre-Crisis Sample
(fixed-effect panel estimates with SUR standard errors in parenthesis; sample: 1995:03-1997:07)[1]

	IND3	*IND2*	*BIS*
γ_1	-1.450 ***	-1.140 ***	-1.632 ***
	(0.328)	(0.227)	(0.427)
γ_2	3.073 ***	2.198 ***	3.185 ***
	(0.862)	(0.642)	(1.098)
$\underline{\gamma}$	7.297 ***	7.572 ***	6.996 ***
	(0.267)	(0.344)	(0.244)
γ_3	-0.008	-0.025	0.029
	(0.024)	(0.018)	(0.029)
ρ	0.375 ***	0.149 *	0.608 ***
	(0.066)	(0.077)	(0.065)
R^2	0.384	0.094	0.392
DW	1.507	1.477	1.577
Observations	174	174	174

[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)[1]

	Thailand	*Indonesia*	*South Korea*	*Malaysia*	*Singapore*	*Hong Kong*
$\gamma_{0,j}$	-9.609	-1.294	-12.448	-14.402 *	-18.941	7.397
	(8.147)	(1.604)	(8.576)	(8.020)	(12.549)	(6.090)
γ_1	-0.290 *					
	(0.159)					
γ_2	0.400 ***					
	(0.085)					
$\underline{\gamma}_j$	*estimated recursively (see Fig. 6)*					
$\gamma_{3,j}$	0.113	0.034 *	0.168 *	0.153 **	0.178	-0.062
	(0.087)	(0.020)	(0.102)	(0.071)	(0.113)	(0.046)
ρ_j	0.385 **	0.250	0.487 *	0.418	0.285	0.154
	(0.159)	(0.182)	0.264	(0.273)	(0.209)	(0.207)
Observations	76	76	76	76	76	76

[1] Data are monthly.
Three (***), two (**), and one (*) stars mark statistical significance respectively at one, five, and ten percent levels.
The coefficients γ_1 and γ_2 are restricted to be the same across countries.

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