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Abstract

This paper studies empirically how uncertainty affects speculation in the foreign exchange markets. We use the dispersion of survey forecasts of key macroeconomic variables to measure uncertainty about fundamentals. We find that uncertainty has a *non-monotone* effect on exchange rate pressures: namely, uncertainty heightens speculative pressures when expected fundamentals are good and eases them when they are bad. We prove that this prediction arises from a broad class of currency crisis theories, ranging from first-generation to global-game models. We also show that the proposed empirical strategy remains valid in the presence of forecasters with strategic objectives and use a novel set of instrumental variables to address potential endogeneity bias.

JEL Classification: F31, D84, D82

Keywords: First-generation models, Global games, Information, Speculation.

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1 Introduction

Recent theoretical contributions to the currency crisis literature have turned the spotlight on the role of uncertainty during crises, showing that the same fundamentals may or may not lead to a speculative attack depending on the precision of information about them. The matter is also critical for policy purposes. For example, if greater uncertainty increases the probability of a speculative attack, then exchange rate regimes will be more vulnerable in periods of high uncertainty and policy-makers should adjust their policies accordingly. There has been little debate, however, about the empirical significance of uncertainty about fundamentals during currency crises. This paper analyzes this question by proceeding in two stages.

First, we show that a broad class of currency crisis theories with a unique equilibrium — ranging from first-generation models with public information (such as those pioneered by Krugman, 1979, and Flood and Garber, 1984) to global games with public and private information (see Morris and Shin, 1998, and Hellwig, 2002) — predicts that the effect of uncertainty on exchange rate pressures is non-monotone and varies with expected fundamentals.[1] Specifically, these models predict that, when expected fundamentals are "good," a reduction in information precision (i.e. an increase in uncertainty) raises the share of speculators attacking the currency, whereas a reduction in information precision with "bad" expected fundamentals has the opposite effect. This broad-based prediction — which previous studies have overlooked — has a very intuitive interpretation: as information about fundamentals becomes less precise, speculators rely less on it in order to decide whether to attack the currency. Thus, as information about *bad* fundamentals becomes less reliable, speculators lose confidence in the success of a speculative attack and diminish exchange rate pressures. By the same token, as information about *good* fundamentals becomes less reliable, speculators lose confidence in the good state of the economy and augment exchange rate pressures.[2]

Second, we study empirically how speculation in the foreign exchange markets is affected by uncertainty about fundamentals, by measuring the latter with the dispersion of survey forecasts of key macroeconomic variables in six Asian countries.[3] Survey forecasts are an appealing source of information to test models in which exchange rate pressures depend on agents beliefs about fundamentals. Since 1995, for the six Asian countries in our

[1] For first-generation models, we derive our results from a framework that nests the models of Grilli (1986) and Goldberg (1991). For global games, we consider the framework analyzed by Hellwig (2002), Metz (2002), and Morris and Shin (2004).

[2] It should be clear that we do not attempt to discriminate between first-generation and global-game models, but we study the role of uncertainty by encompassing different models. On the former matter, our results suggest, if anything, that in order to discriminate between different currency crisis models, one cannot look at their implications about the role of uncertainty, which are essentially the same across models.

[3] The countries in our sample are: Hong Kong, Indonesia, Malaysia, Singapore, South Korea, and Thailand. Data are monthly and cover the period from January 1995 to April 2005.

sample, Consensus Economics has gathered individual forecasts from a panel of analysts including both international financial firms and domestic enterprises at a monthly frequency — a higher frequency than that of some key macroeconomic variables such as GDP growth. Survey forecasts are also inherently forward looking, just like the exchange rate pressures that the currency crisis literature tries to explain. Moreover, it is possible to relate the mean and variance of the individual forecasts to, respectively, expected fundamentals and the precision of information, which are the key parameters of currency crisis models.

The main challenge we face in studying empirically the effect of uncertainty on exchange rate pressures is that causality could go both ways. While currency crisis theories predict a causal effect from the mean and variance of the forecasts to exchange rate pressures, shocks to unobservable determinants of exchange rate pressures can also affect the distribution of the forecasts.

We tease out this complex interaction by building a novel set of instrumental variables. Our instruments include the level and dispersion of forecasts of consumption growth and the unemployment rate in the United States, as well as alternative composite instruments computed as the trade-weighted version of the same macroeconomic indicators for G-6 countries (i.e., the G-7 countries without Japan).

Are these valid instruments for GDP growth forecasts of Asian countries? The ideal instrument is an external source of variation that randomly changes the mean and variance of the GDP growth forecasts of Asian countries, in a manner that is uncorrelated with the unobservable determinants of exchange rate pressures. Forecasts of domestic demand conditions in the United States (or in G-6 countries) satisfy the requirement of being positively correlated with GDP growth forecasts of Asian countries, because the export-oriented economies of the latter depend on cyclical conditions in the former. Moreover, our proposed instruments are exogenous to speculative pressures in Asian countries since they reflect mostly domestic factors in the United States, on which exchange rate developments in Asian countries have a negligible effect.

These two characteristics, however, are not sufficient to fully justify the use of forecasts of domestic demand conditions in the United States as instruments for GDP growth forecasts of Asian countries. The additional identifying assumption which is required is that our instruments have no reason to be included as regressors in the second stage of the IV estimation.[4] In other words, we need to assume that forecasts of domestic demand conditions in the United States only affect exchange rate pressures in Asian countries *indirectly*, i.e. by modifying these countries' GDP growth forecasts. We justify this exclusion restriction by including forecasts of U.S. interest rates, as well as their interaction with lagged levels of

[4]This exclusion restriction is critical because, if domestic demand conditions in G-6 countries needed to be included in the second stage and we omitted them, our instruments would be correlated with the residual unobservable determinant of exchange rate pressures.

short-term debt in each Asian country. These variables aim at capturing changes in financial market conditions in the United States that might affect exchange rate pressures in Asian countries — for example, through higher expected costs of servicing their external debt — and that our instruments might proxy. Then, our identifying assumption is that, after controlling for U.S. interest rates and their interaction with Asian countries' short-term external debt, forecasts of domestic demand conditions in the United States do not have any residual *direct* effect on Asian countries' exchange rate pressures.

Using IV estimates is very important also to address another problem that could potentially affect our analysis. The mean forecasts is only an imperfect measure of expected fundamentals (think, for instance, at the sample error due to the fact that we observe only a finite number of forecasters). Therefore, in a multivariate specification like the one that is predicted by the theory, OLS would provide biased estimates of all coefficients (in a direction that is difficult to assess), including also the coefficient related to uncertainty, which is the main variable we are interested in. Instrumental variables, then, are our main weapon to address this second problem as well.

Table 1 presents some preliminary evidence about the effect of uncertainty on exchange rate pressures. The north-west panel of the table shows that, when GDP growth forecasts are bad, high uncertainty is associated with low exchange rate pressures.[5] Conversely, the north-east panel of the table shows that, when GDP growth forecasts are good, high uncertainty is associated with high exchange rate pressures. Both findings are consistent with the *non-monotone* effect predicted by the theory that uncertainty heightens speculative pressures when expected fundamentals are good and eases them when they are bad. These results, which are confirmed by simple OLS regressions, are corroborated by our IV estimates, as well as by an extensive set of robustness checks.

There are relatively few studies in the empirical literature on currency crises that have focused on the role of uncertainty. Early 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. Our paper differs from these studies for the novel testable prediction derived from a broad class of currency crisis theories with a unique equilibrium and the use of survey data.

[5] In Table 1, the thresholds separating good from bad expected fundamentals and high from low uncertainty are, respectively, the median value of GDP growth forecasts and the median value of the variance of GDP growth forecasts. In our regressions, however, the threshold separating good from bad fundamentals is estimated (while no threshold is needed for the variance). Following Eichengreen, Rose, and Wyplosz (1996), our measure of exchange rate pressures is a weighted average of changes in the exchange rate, changes in international reserves which can be paid out in responding to pressures, and changes in the domestic interest rates (since interest rates can be raised to fend off an attack). Section 3 explains in detail the construction of this variable. We thank Jonathan Eaton for the suggestion of including this table in the paper.

More recently, Bloom, Floetotto, and Jaimovich (2009) have documented that uncertainty, as measured by the volatility of several micro and macro economic variables, is countercyclical. Interestingly, these findings, coupled with ours, imply that changes in uncertainty contribute to make exchange rate pressures *less* cyclical. Thus, an intriguing result from this analysis is that, for what concerns speculative pressures, the effect of uncertainty is not as bad as it is generally thought.

The rest of the paper is organized as follows. Section 2 presents the testable implications on the role of uncertainty and expected fundamentals that we derive from both first-generation and global-game models. Section 3 describes the data. Section 4 illustrates the empirical results. Section 5 concludes.

2 Theoretical underpinnings

To analyze the role of uncertainty during currency crises, we estimate variants of the following equation:

$$ERP = \gamma_0 + \gamma_1 y + \gamma_2 \sigma (y - \underline{\gamma}) + \varepsilon \ , \tag{1}$$

where ERP is an index of the speculative pressures, y represents the expected fundamentals, σ is a measure of uncertainty, $\underline{\gamma}$ is a threshold separating good from bad expected fundamentals, and ε is a residual. The most important feature of equation (1) is that it allows for a non-monotonic effect of uncertainty: this effect is, in fact, assumed to depend on the level of expected fundamentals. Hence, for instance, if $\gamma_2 > 0$ (as the theory predicts), higher uncertainty will augment pressures on the exchange rate only for sufficiently good expectations ($y > \underline{\gamma}$). Conversely, higher uncertainty will attenuate exchange rate pressures with bad expectations ($y < \underline{\gamma}$).

In principle, one could think that higher uncertainty always encourages speculative attacks.[6] The purpose of this section is to show, instead, that a non-monotonic effect of uncertainty emerges in both first-generation models with public information and global games with public and private information. These models, which share the key feature of yielding a unique equilibrium, allow us to derive exact comparative statics results and to bring to light the effects of expected fundamentals and uncertainty reflected in equation (1).[7]

[6] For example, uncertainty has a negative effect on aggregate output and employment in the micro model of the firm of Bloom (2009) as well as in the dynamic stochastic general equilibrium macro model of Bloom, Floetotto, and Jaimovich (2009). In the currency crisis literature, Flood and Marion (2000) build a model with multiple equilibria in which higher uncertainty about a real shock raises the risk premium on asset holdings, then expected fundamentals deteriorate, moving the economy to a region where a self-fulfilling attack may occur. Similarly, uncertainty seems to have a negative effect in Flood and Garber (1984), because their model assumes an exponential distribution of fundamentals in which the effect of higher uncertainty cannot be distinguished from the effect of deteriorating expected fundamentals.

[7] Second generation models of currency crisis à la Obstfeld (1996) are not suitable to analyze the role of

2.1 Krugman-Flood-Garber models

Following Krugman (1979) and Flood and Garber (1984), first-generation models consider dynamic settings with deteriorating fundamentals, usually represented by excessively increasing money supply. These models build on three key equations: the uncovered interest rates parity (UIP), the purchasing power parity (PPP), and the money market equilibrium condition (MME) — an equation in which, in its simplest form, real money supply is a function of nominal interest rates (see Appendix A for analytic details).

The exchange rate, defined in terms of units of domestic currency per one unit of foreign currency, is fixed by public authorities at a peg $\overline{e}$. A speculative attack is assumed to occur *as soon as* speculators can earn positive expected profits from a devaluation (thus, neglecting possible coordination problems, which are the focus of other models). Positive expected profits, in turn, are obtained whenever the expected shadow exchange rate — defined as the exchange rate that would prevail after the devaluation — becomes larger than the peg. Hence, at time t, the probability of a speculative attack is $\Pr\left(\widetilde{e}_{t+1} > \overline{e}\right)$, where $\widetilde{e}_{t+1}$ is the expected shadow exchange rate at time $t+1$. In Appendix A we show that, taking into account the three key equations, this probability translates into:

$$\Pr\left(\Theta_{t+1} \leq k_t\right) \ , \tag{2}$$

where Θ_{t+1} is random variable representing the stochastic component of the fundamentals at time $t+1$, and $k_t \in \mathbb{R}$ gathers all the constants and the components of the fundamentals whose values are known at time t. The signs of the variables are chosen so that a decrease in Θ_{t+1} and an increase in k_t correspond to deteriorating fundamentals.

Various first-generation models differ only in the specification of Θ_{t+1} and k_t. In Flood and Garber (1984), Θ_{t+1} is a stochastic shock to the money supply at time $t+1$; k_t is a function of the time-t levels of international interest rates, foreign prices, the exchange-rate peg, the domestic credit, its drift, and the lower bound for international reserves (usually set to 0). Similarly, Grilli (1986) gathers all the fundamentals into a variable h_t, which follows an $AR\,(1)$ process with some positive drift; then, Θ_{t+1} is the stochastic shock of this process at time $t+1$, while k_t includes the variables considered by Flood and Garber (1984), plus the sensitivity of money demand to the exogenous real domestic income.[8]

uncertainty because they typically assume complete information and they yield multiple equilibria, reducing the possibilities to derive precise comparative statics results. Sbracia and Zaghini (2001) and Prati and Sbracia (2010), however, consider incomplete information extensions of these models and show that they yield predictions about the space of parameters in which a speculative attack is feasible that are consistent with the predictions analyzed in this section.

[8] Goldberg (1991) generalizes all three key equations: the UIP, by considering a risk premium; the PPP, by introducing non-traded goods, and systematic and random deviations from PPP; the MME, by adding a currency substitution motive to money demand. In her model, Θ_{t+1} is a linear function of three independent random variables: the stochastic deviations from PPP, the shocks to domestic credit (to finance unanticipated government expenditure), and the uncertain availability of external credit.

As in Grilli (1986), we assume that $\Theta_{t+1} \sim Norm(y, 1/\alpha)$, i.e. that Θ_{t+1} is normally distributed with mean $y \in \mathbb{R}$ and variance $1/\alpha$, where $\alpha > 0$.[9] Since this distribution is common knowledge, we can regard it as the *public information* available to speculators. Thus, we can refer to the expected fundamental y as the "public signal" and to α as the "precision of public information."

The probability of a speculative attack can be written as $\Phi\left[\sqrt{\alpha}\,(k_t - y)\right]$, where Φ is the cumulative distribution function (cdf) of the standard normal distribution. By differentiating, it is immediate to show the following result:

Proposition 1 *The probability of a speculative attack,* $\Pr(\Theta_{t+1} \leq k_t)$*, is: (i) decreasing in* y*; (ii) decreasing (increasing) in* α *if* $y > k_t$ $(y < k_t)$*.*

The proposition shows that the probability of a speculative attack is always reduced by an improvement in expected fundamentals. The most striking result, however, is that the effect of the precision of public information depends on the level of expected fundamentals. Specifically, if y is sufficiently good (i.e. $y > k_t$), then an increase in α diminishes the probability of a currency crisis, and vice versa if y is sufficiently bad $(y < k_t)$. Before providing some intuition, we discuss the extension of this result to other distributions and, in the next section, we show that this prediction is also shared by another class of currency crisis models.

Throughout the paper, we use the normality assumption for two main reasons. First, to maintain consistency across models, as global games with public and private information can be easily solved with normal distributions. Second, because when we test the normality assumption of the survey forecasts that we use in the empirical analysis, we cannot reject it at the standard 5 percent threshold. Nonetheless, the possibility of collapsing first-generation models into an equation as simple as (2) allows for a straightforward generalization of the main results to non-normal distributions. Assume that Θ_{t+1} is a random variable with a location parameter l and a scale parameter s, and denote its cdf by $F(\cdot; l, s)$. The definitions of l and s (which, in the normal distribution, correspond to the mean and the standard deviation) imply that: $F(\theta; l, s) = F\left(\frac{\theta - l}{s}; 0, 1\right)$. Note that the scale parameter is the inverse of the precision parameter. Hence, we can set $1/s = \rho$, where ρ denotes the precision, so that the probability of a speculative attack becomes: $\Pr(\Theta_{t+1} < k_t; l, \rho) = F(\rho(k_t - l); 0, 1)$. By the properties of distribution functions, it immediately follows that the probability of a speculative attack is decreasing in the location parameter l, and decreasing (increasing) in the precision parameter ρ if $l > k_t$ $(l < k_t)$, generalizing Proposition 1.

[9] First-generation models usually set $y = 0$, as they gather into k_t all the constants (including the drift of the stochastic process driving the fundamentals).

2.2 Morris-Shin models

Morris and Shin (1998, 2004) and Hellwig (2002) modify the previous framework along three main dimensions. First, following Obstfeld (1996), they introduce a coordination problem: for some range of fundamentals, only a coordinated attack can force the government to abandon the peg; as a result, complete information about fundamentals would yield multiple equilibria. Second, they show that an incomplete-information version of the model with public and private information about fundamentals yields a unique equilibrium if private signals are sufficiently precise relative to public signals. Third, in order to focus on coordination and information issues, they replace the dynamic setting with a simpler static one.

To model strategic complementarities, Morris and Shin assume that public authorities devalue the currency whenever the share of attackers, $l \in [0, 1]$, is larger than the fundamental θ. Payoffs are as follow: agents who successfully attack the currency peg obtain a net payoff $D - \tau > 0$, where $\tau > 0$ is a transaction cost; unsuccessful attacks pay the cost $-\tau$; refraining from attacking provides 0. This complete-information problem yields a tripartition of the space of fundamentals: if $\theta \in (-\infty, 0]$, the government abandons the peg; if $\theta \in (1, +\infty)$, the government maintains the peg; if $\theta \in (0, 1]$, there are multiple equilibria: speculators can either attack the currency and force a devaluation, or refrain from attacking and allow the peg to be maintained.

Now assume that agents do not observe θ before deciding whether to attack and, as in the previous section, take a normal distribution of the fundamentals, $\Theta \sim Norm(y, 1/\alpha)$. Assume also that each speculator i receives a *private signal* $x_i = \theta + \varepsilon_i$, with $\varepsilon_i \sim Norm(0, 1/\beta)$ (with ε_i and ε_j independent given θ for each $i \neq j$), where $\beta > 0$ is the precision of private information. It is possible to show that this model yields a unique equilibrium if $\beta \geq \alpha^2/2\pi$ (Morris and Shin, 2004). This equilibrium consists of a unique value of the private signal x^* such that each speculator receiving a signal lower than x^* attacks the currency peg (trigger strategy), and a unique level of the fundamentals $\theta^* \in [0, 1]$ such that the government abandons the peg when fundamentals are lower than θ^*.[10]

The equilibrium trigger points θ^* and x^* are both functions of y, α, and β. Thus, speculators' expectations matter and one can derive rigorous comparative statics about their effects on the trigger points θ^* and x^*. Most importantly, it is possible to calculate the effects of the parameters on the probability that speculator i attacks, $\Pr(X_i \leq x^* \mid \theta)$. This probability represents *the share of speculators attacking the currency* and, therefore, has a straightforward empirical counterpart in an index of exchange rate pressure.

[10] The original Morris-Shin game has been subject to several extensions, encompassing other distributions (Heinemann and Illing, 2002) and "large" speculators (Corsetti, Dasgupta, Morris, and Shin, 2004). Many studies (such as Angeletos, Hellwig, and Pavan, 2003, and Tarashev, 2007) have focused on modifications of the model that restore multiplicity of equilibria, such as the introduction of signalling by central banks and the information-aggregation role of interest rates.

The share of speculators attacking the currency is equal to $\Phi\left[\sqrt{\beta}\,(x^* - \theta)\right]$. This probability depends on the actual fundamental θ and on the parameters y and α (which affect x^*) and β (which is in the argument of Φ and also affects x^*). In Appendix B, we show the following result:

Proposition 2 *Assume that $\beta \geq \frac{\alpha^2}{2\pi}$; then the probability $\Pr(X_i \leq x^* \mid \theta)$ is: (i) decreasing in y; (ii) decreasing (increasing) in α if $y > s_1$ ($y < s_1$), where $s_1 \in \mathbb{R}$; (iii) decreasing (increasing) in β if $\theta > s_2$ ($\theta < s_2$), where s_2 is a function of y.*[11]

As in first-generation models, an improvement in expected fundamentals always reduces the share of speculators attacking the currency (point *(i)* of Proposition 2). Similarly, the effect of the precision of public information depends on expected fundamentals (point *(ii)*): when y is sufficiently good, an increase in α strengthens the belief that fundamentals are good and reduces the share of attacking speculators by decreasing the equilibrium trigger point x^* (and vice versa when y is bad).

The effect of changes in the precision of private information β on the share of attackers is more complex. This effect depends not only on the expected fundamental y but also on the realized fundamental θ, around which the signals x_i are centered. Two forces are at work. First, there is an effect of β which is opposite to the effect of α, due to the fact that the ex-post expected fundamental of agent i is a weighted average of y and x_i with weights respectively equal to α and β. If y is high, speculators expect good fundamentals. Then, if the precision ratio α/β is high (low), speculators know that also the other speculators have formed their expectations attributing a large (small) weight to the good public signal; then, they will be less (more) inclined to attack the currency. In other words, coordination on a good public signal is easier (harder) when the public component y in each individual expectation carries a large (small) weight. On the contrary, for a low y, a low ratio α/β tends to reduce the share of attackers because agents regard that bad public signal as less reliable; by the same token, a high ratio α/β raises the share of attackers for the opposite reason.

A higher β, however, has also the additional effect of concentrating the private signals around the actual fundamental θ. This effect can either offset or reinforce the first one, depending on the realization of θ. Suppose y is good. If the realized θ is also good, the number of signals below the threshold x^* may diminish enough to offset the effect of a higher x^* (due to the increase in β when y is good); as a result, the share of attackers decreases following the increase in β. In this case, β has the same effect as α. If the realized fundamental θ is, instead, bad, the second effect reinforces the first and the share of attacking speculators increases; i.e. β and α have opposite effects. When we consider the polar case in which y is bad, we have symmetric results. In particular, if the realized fundamental θ is also bad,

[11] Exact expressions of s_1 and s_2 are in Appendix B.

the effect of the greater concentration of the private signals around θ may dominate and the share of attackers increases, as in the case of an increase in α.

Thus, changes in the precision of private information tend to have the same effect as changes in the precision of public information, provided that the average private signal (i.e., the actual fundamental) and the public signal (the ex-ante expected fundamental) are either both sufficiently good or both sufficiently bad. Instead, if the offsetting effect of actual fundamentals is not sufficiently strong, then the precision of private information has effects that are opposite to those of public information.[12]

2.3 Testable implications

What are the empirical implications of the models discussed above for the signs of γ_1 and γ_2 in equation (1)? Clearly, both first-generation and global-game models predict that an improvement in expected fundamentals, y, eases the pressures on the exchange rate. Therefore, we expect γ_1 to be negative.

Now suppose that the measure of uncertainty σ reflects the inverse of the precisions of public and private information (as shown in Appendix C). If α increases (i.e. σ decreases) with good expected fundamentals ($y > \underline{\gamma}$), then both models predict that exchange rate pressures decline. Therefore, theoretical models imply $\gamma_2 > 0$. Similarly, if α increases with bad expected fundamentals ($y < \underline{\gamma}$), then exchange rate pressures rise. Hence, we still expect $\gamma_2 > 0$. Obviously, in the mirror cases where α diminishes, the implications for exchange rate pressures are opposite, so that the effect of α always implies $\gamma_2 > 0$.

The effect of an increase in β can be derived from the global-game model. If β changes and actual and expected fundamentals are both sufficiently good or both sufficiently bad, then the effect of β is the same as the effect of α, thereby implying $\gamma_2 > 0$. Hence, while there are very good reasons to expect $\gamma_2 > 0$, we cannot completely rule out the possibility of a negative sign, because we do not have separate observations for α and β. Thus, it might be $\gamma_2 < 0$ if a change in σ is due to a change in β with a low y and a high θ, or vice versa.

[12] While a generalization of the Morris-Shin model is beyond the scope of this paper, we can offer some hints on the possibility of extending Proposition 2 to other distributions. Consider a variable X_i with location parameter l (the actual fundamental with the normality assumption), precision parameter ρ (the precision of private information), and cdf $F(\cdot; l, \rho)$ and assume that the model yields a unique equilibrium trigger point x^*, so that $F(x^*; l, \rho)$ is the share of attackers. In the equilibrium, x^* changes not only with the precision of private information ρ, but also with expected fundamentals y and the precision of public information α (y and α being the parameters of the prior). If, as it is reasonable, $\partial x^*/\partial y < 0$ (i.e. the trigger point for attacking the currency decreases as expected fundamentals improve), then part (i) of Proposition 2 holds. Analogously, note that $\partial F(x^*; l, s)/\partial\alpha$ has the same sign as $\partial x^*/\partial\alpha$, while $\partial F(x^*; l, s)/\partial\rho$ has the same sign as $\left[(x^* - l) + \rho\frac{\partial x^*}{\partial\rho}\right]$. These results, when compared with equations (12) and (13), suggest that the same non-monotone effects of uncertainty found for normal distributions also hold for other distributions.

In these situations, the expected sign depends on which of the two effects described above prevails (namely, that of α/β as relative weight of y versus that of $1/\beta$ as dispersion of the private signals around θ).

3 The data

To estimate variants of equation (1), we need to measure exchange rate pressures, find the empirical counterparts of y and σ, build the necessary instruments, and include other controls that are typically considered in empirical studies of currency crises. In this section, we describe all these variables.

3.1 Indices of speculative pressure

To measure the fraction of speculators that decide to attack the currency, we build an index of exchange rate pressure.[13] Previous empirical studies have generally used continuous indices of exchange rate pressure as an intermediate step towards the construction of a binary zero-one variable corresponding to tranquil and crisis periods. In this paper, we use the underlying continuous index of exchange rate pressure directly as a dependent variable. The aim is to have an empirical approach that encompasses the private information models considered in this paper, where some speculators attack the currency while others do not, so that the number of attackers varies continuously with fundamentals and agents' beliefs. This approach makes sense as, in reality, different groups of agents may take opposite positions against any given currency.

Our index of exchange rate pressure $IND3$ is the sum of the normalized values of three indicators:[14] (*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.[15]

[13] Eichengreen, Rose, and Wyplosz (1996) first developed empirical indices of exchange rate pressures, whose theoretical foundations lie in monetary models of optimal exchange-rate-market intervention.

[14] To normalize, we subtract from each indicator the country-specific mean and divide the result by the country-specific standard deviation.

[15] To check the robustness of our results, we also compute two other indexes. An index $IND2$, which sums only normalized values of the indicators (*i*) and (*ii*) and a second index BIS, which is the continuous version of an index developed by the Bank for International Settlements also based on changes in exchange rates, international reserves, and real interest rates (see Hawkins and Klau, 2000). In the rest of the paper, we present estimation results for $IND3$; those for $IND2$ and BIS, which are very similar, are reported in Prati and Sbracia (2002).

3.2 Survey forecasts of domestic macroeconomic variables

Every month, Consensus Economics gathers forecasts of a series of macroeconomic 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 in 1997.

To reproduce a constant forecast horizon of one year, we follow Brooks, Edison, Kumar, and Sløk (2004) and 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.[16] Appendix C shows the link between the mean forecast and expected fundamentals, and between the variance of the forecasts and the precision of public and private information. 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.[17]

3.3 Instrumental variables

To address possible endogeneity bias, we create a novel set of instrumental variables. These are the level and dispersion of forecasts of consumption growth and the unemployment rate in the United States, as well as an alternative composite instrument computed as the trade-weighted version of the same variables for G-6 countries (i.e., the G-7 countries without Japan). We use forecasts of consumption growth and the unemployment rate instead of GDP growth forecasts to make our instruments as exogenous as possible to crisis events in Asian countries, given that the export component of G-6 GDP might be partially affected by the

[16] 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 each year than at the beginning of each year.

[17] Using the mean instead of the median forecast or using the standard deviation instead of the mean absolute median difference would provide essentially the same results. Consider that correlation between the mean and the median forecast for our six countries is between 99.2 and 99.5 percent, while that between the standard deviation and the mean absolute median difference is only slightly lower. Figures 1 and 2 in Prati and Sbracia (2002) show all the four variables for the six countries in our sample.

economic crisis in Asia.[18] Excluding Japan from the composite instrument is an additional precaution aimed at ensuring its exogeneity in view of the somewhat greater dependence of Japan's economy on developments in neighboring Asia. Finally, the trade weights used to construct the composite index are based on pre-crisis export shares over the period 1985-1994.

3.4 Other controls

In addition to survey forecasts of macroeconomic variables, we tried other regressors either used in previous empirical studies on currency crises or suggested by the theoretical literature (as reported also in Appendix D). Only three additional variables — the real exchange rate, a forecast of U.S. interest rates, and the interaction between such forecast and BIS short-term external debt — had a robust and significant impact across specifications and, therefore, were included in the regressions.

The real effective exchange (computed by JP Morgan Chase) had a better overall fit then the nominal exchange rate vis-à-vis the U.S. dollar. There was no difference between the two models, however, in terms of the estimated signs and significance of all other coefficients.

Consensus Economics provides 3-month- and 12-month-ahead forecasts of 3-month U.S. interest rates and 10-year U.S. treasury bond yields. Given that these four forecasts are highly correlated, we tried each of them separately and selected the most significant one (the 3-month ahead forecast of 3-month rates). These forecasts were also superior to current U.S. interest rates of various maturities, including the federal funds rate.

We capture the vulnerability associated with the combination of high U.S. interest rates and high short-term external debt of most Asian countries with the interaction between our preferred U.S. interest rate forecast and the ratio of BIS short-term external debt to population (expressed in deviation from its country mean). This variable reflects the non-linear effect of U.S. interest rates on speculative pressures, which is an increasing function of short-term debt. Expressing short-term debt as a ratio to population rather than GDP gives us a measure of indebtedness that fluctuates mostly with short-term debt rather than with GDP, which was very volatile during the crisis period. For the same reason, we do not scale short-term debt with international reserves.

[18] Nonetheless, the small weight of exports in U.S. GDP combined with the small share of U.S. exports to the six Asian countries in our study would make U.S. or G-6 GDP growth forecasts an instrument reasonably exogenous to exchange rate pressures in Asian countries. IV regressions with U.S. or G-6 GDP-growth forecasts used as instruments yield results similar to those we obtained using instruments based on consumption and unemployment rate.

4 Estimation results

This section presents estimation results for a simple baseline regression, our key instrumental variable regressions, and a pre-crisis panel model. Appendix D reports additional robustness tests based on a dynamic model (including a panel GMM specification), as well as models with time-varying coefficients and time-varying thresholds separating good from bad fundamentals.

4.1 Baseline regression

We estimate the following version of equation (1):

$$\begin{aligned} IND3_{j,t} &= \hat{\gamma}_{0,j} + \hat{\gamma}_1 f^e_{GDP_j,t-1} + \hat{\gamma}_2 \sigma^e_{GDP_j,t-1} \cdot (f^e_{GDP_j,t-1} - \underline{\hat{\gamma}_j}_{GDP}) \\ &+ \hat{\gamma}_3 i_{US,t-1} + \hat{\gamma}_4 i_{US,t-1} \cdot ST_{j,t-1} + \hat{\gamma}_{5,j} e_{j,t-1} + u_{j,t} \ , \\ &\text{with} \quad u_{j,t} = \hat{\rho}_j u_{j,t-1} + \varepsilon_{j,t} \end{aligned} \tag{3}$$

where $IND3_{j,t}$ is our three-component index of exchange rate pressure for country j at time t, $f^e_{GDP_j,t-1}$ is the median forecast of GDP growth, $\sigma^e_{GDP_j,t-1}$ is the mean absolute median difference of GDP growth forecasts, $i_{US,t-1}$ is the 3-month ahead forecast of 3-month U.S. interest rates, $ST_{j,t-1}$ is the ratio of short-term external debt to population of country j at time $t-1$ expressed in deviation from country j's average ratio over the sample, and $e_{j,t-1}$ is the real effective exchange rate.

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 $\hat{\gamma}_1$, $\hat{\gamma}_2$, $\hat{\gamma}_3$, and $\hat{\gamma}_4$, could be constrained to be the same across countries (the null hypothesis of equality was not rejected with a p-value of 0.345). The first six columns of Table 2 shows the results of this restricted estimation of (3). We use the restrictions accepted by the data to simplify the presentation and to conduct coherent robustness tests involving recursive estimation (see Appendix D) on a specification with a small number of parameters. The restriction is by no means necessary to obtain statistically significant coefficients: in the unrestricted estimates, all $\hat{\gamma}_{1,j}$ were negative and statistically significant at the 5 percent confidence level and all $\hat{\gamma}_{2,j}$ were positive and statistically significant at the 1 percent confidence level.

The results reported in the first six columns of Table 2 confirm that higher expected GDP growth reduces exchange rate pressures ($\hat{\gamma}_1 < 0$). Most interestingly, these estimates indicate that uncertainty about GDP growth has an additional effect, which depends on expected GDP growth, as the theory predicts. A higher dispersion of GDP growth forecasts increases exchange rate pressures when expected GDP growth is above the estimated country-specific threshold and reduces them when it is below ($\hat{\gamma}_2 > 0$). Thresholds are statistically different from zero only for Thailand (equal to 2.5 percent) and Malaysia (4 percent).

Two other results also stand out from this analysis. Uncertainty about GDP growth appears to be the most important variable in our regression and provides the largest contribution to the overall goodness-of-fit of the model when it is interacted with expected GDP growth, while when it is not interacted it remains significant but its contribution to the overall goodness-of-fit is much smaller. In particular, the contribution of the interaction term to the overall goodness-of-fit measure is larger than that of the median forecast alone: if we exclude $f^e_{GDP_j,t-1}$ by setting $\hat{\gamma}_1 = 0$ in (3), the R^2 for the overall system falls from 42.6 to 37.6 percent, whereas if we exclude $\sigma^e_{GDP_j,t-1} \cdot (f^e_{GDP_j,t-1} - \underline{\hat{\gamma}_j}_{GDP})$ the R^2 for the overall system drops from 42.6 to 33 percent. Adding back the dispersion of the forecasts $\sigma^e_{GDP_j,t-1}$ to the list of regressors without interacting it with $(f^e_{GDP_j,t-1} - \underline{\hat{\gamma}_j}_{GDP})$ yields an R^2 of only 34.3 percent.

The interaction between U.S. interest rates and short-term debt is strongly significant suggesting that global financial conditions have an impact on speculative pressures through the level of indebtedness of each country.[19] An appreciated real exchange rate is also associated with higher speculative pressures in all countries except Hong Kong.

In this regression we have to worry about reverse causality from exchange rate pressures to the distribution of GDP growth forecasts. In particular, one could easily think of a model in which higher exchange rate pressures raise uncertainty about GDP growth forecasts.[20] On the other hand, it is difficult to think of a model in which higher exchange rate pressures lower uncertainty and, even more so, of a model in which higher exchange rate pressures lower uncertainty when expected fundamentals are bad and raise them when expected fundamentals are good. If causality went from exchange rate pressures to uncertainty, why should this effect depend on the level of GDP growth forecasts?[21] While there are good reasons to think that reverse causality is not the main driver of the previous results, in the next section we address the causality issue directly, using instrumental variables. This method also allows us to address possible omitted variable problems and the problem connected to the fact, explained in Appendix C, that the mean and variance of the forecasts are imperfect measures

[19] Prati and Sbracia (2002) estimate also a "minimal" regression without additional controls like U.S. interest rates and their interaction with short-term debt and show that $\hat{\gamma}_1$ and $\hat{\gamma}_2$ are still significant and with the predicted sign.

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

[21] The focus on the theoretical mechanism through which uncertainty should affect exchange rate pressures as a way to make progress on the causality issue is akin to the approach of Rajan and Zingales (1998). These authors use the theoretical prediction that financial development should help disproportionately firms that are more dependent on external finance for growth to shed light on the causal effect of financial development on growth.

of expected fundamentals and uncertainty about fundamentals.

4.2 Instrumental variable estimates

Lagging the mean and variance of the forecasts by one month — as we do in the baseline regression — rules out *contemporaneous* reverse causality. Predetermined mean and variance of the forecasts could, however, still be endogenous if a *serially correlated* omitted variable, including a possible measurement error, affected exchange rate pressures and the distribution of the forecasts. In this case, lagged values of the mean and variance of the forecasts and the error term would not be independent. We address this potential problem by instrumenting $f^e_{GDP_j,t-1}$ and $\sigma^e_{GDP_j,t-1}$ with the median and absolute median difference of the Consensus forecasts of two macroeconomic variables, consumption growth and unemployment rate, capturing cyclical domestic demand conditions in the United States and in G-6 countries.

The rationale for these instruments is threefold: (*i*) they are positively correlated with GDP growth forecasts for Asian countries, whose export-oriented economies depend on cyclical developments in G-6 countries; (*ii*) they are exogenous to speculative pressures in Asian countries because they reflect mostly domestic demand developments in G-6 countries; and (*iii*) they can be excluded from the second-stage regressions because they are likely to affect exchange rate pressures in Asian countries only indirectly through Asian countries' GDP growth forecasts. The third exclusion restriction requirement is satisfied on the grounds that the presence among the regressors of the forecasts of U.S. interest rates and their interaction with Asian countries' short-term debt captures the impact of global financial conditions on speculative pressures. In other words, if we did not include these variables, our instruments might have proxied them in view of their likely correlation with financial conditions in the United States through the Federal Reserve's reaction function. For example, forecasters might expect the Federal Reserve to raise U.S. interest rates in response to a forecasted tightening of the U.S. labor market.

The first six columns of Tables 3 and 4 show the results of the regressions with the first two instruments, the forecasts of consumption growth and of unemployment rate in the United States. The estimated coefficients $\hat{\gamma}_1$ and $\hat{\gamma}_2$, which are the focus of our study, remain statistically significant and with the expected sign. Results are broadly unchanged, however, also for what concerns the other coefficients. In particular, $\hat{\gamma}_4$ and $\hat{\gamma}_{5,j}$ preserve the same sign and similar p-values, while $\hat{\gamma}_3$, which was not significant in the baseline regression, becomes weakly significant when the instruments are based on the forecasts of consumption growth and strongly significant when we use the forecasts of the unemployment rate.

Similarly, the first six columns of Tables 5 and 6 confirm that all the coefficients remain largely unchanged when we instrument $f^e_{GDP_j,t-1}$ and $\sigma^e_{GDP_j,t-1}$ with the median and absolute median difference of the forecasts of consumption growth and unemployment rate in G-6 countries. In particular, $\hat{\gamma}_1$ and $\hat{\gamma}_2$ are statistically significant and with the expected sign across all 3SLS estimates.

4.3 Pre-crisis panel model

Should we test the model predictions only on the pre-crisis sample? *Prima facie*, this would seem the correct approach, because in the theoretical models the government pegs the exchange rate. Yet, there are institutional, empirical, and theoretical reasons why we should test the model predictions on the entire sample.

From an institutional point of view, Hong Kong (which maintained its currency board) and Singapore did not formally change their exchange rate regime during the crisis; in other words, on a *de jure* basis, these countries never experienced an exchange rate crisis.[22] On the other hand, some countries that experienced a crisis on a *de jure* basis, like Indonesia, South Korea, and Thailand, adopted a managed float rather than a free float as a post-crisis regime; thus, they adopted an exchange rate regime whose features can still be captured by the models of Section 2. Similarly, Malaysia repegged the exchange rate as early as September 1998, so that the period in which the currency floated was very short.

From an empirical point of view, countries that abandoned pegs battled for avoiding large depreciations of their currencies well beyond July 1997 (the initial date of the crisis). In fact, for all countries but South Korea, the largest outflows of international reserves were recorded in the second half of 1997, when they had already abandoned their pegs. In some cases, the depletion of official reserves continued in the first quarter of 1998 and recurred after the Russian crisis. These are signals that these countries tried to prevent a depreciation of their currencies even after the collapse of the pre-crisis exchange rate regimes.

From a theoretical point of view, even in a floating exchange rate regime speculators 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 the second-generation model analyzed in Section 2.2.[23]

These considerations suggest that the full-sample estimates of the first six columns of Tables 2 to 6 represent a meaningful test of the model. Nonetheless, it is still interesting to

[22] During the crisis, however, Singapore claimed to have broadened the undisclosed target band within which the Singapore dollar was allowed to fluctuate.

[23] 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 would produce downward or upward pressures on the currency through the same mechanism discussed for the other models considered in this paper.

verify whether our results would change if we restricted the analysis to the pre-crisis period 1995:01-1997:07. The seventh column in Tables 2 to 6 shows the outcome of this exercise. Because of the substantial reduction in the number of observations in this sample, we restrict also $\underline{\hat{\gamma}_j}$, $\hat{\gamma}_{5,j}$, and $\hat{\rho}_j$ to be the same across 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 non-monotone 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 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.

We further checked the robustness of our results by re-estimating the full-sample model of the first six columns of Tables 2 to 6 with a set of step dummies set to 1 when a country no longer pegged its exchange rate. The results were essentially unchanged, with $\hat{\gamma}_1$ and $\hat{\gamma}_2$ remaining very significant. Nor did the results change when the pre-crisis panel model 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 $\hat{\gamma}_2$ (discussed in Appendix 3) provides another indication that our results also hold in the pre-crisis sample.

5 Conclusions

Does uncertainty heighten or ease speculation in foreign exchange markets? Could the effect of uncertainty have different signs depending on the economic outlook? To analyze these questions, we develop an empirical framework based on a broad class of currency crisis theories with a unique equilibrium. As a first contribution, we prove that both first-generation models with public information and global games with public and private information predict that uncertainty has a non-monotone effect on speculative pressures; namely, uncertainty heightens speculative pressures when expected fundamentals are good and eases them when they are bad. By measuring expected fundamentals and uncertainty with the mean and the dispersion of survey forecasts, we apply the proposed empirical framework to six Asian countries, finding strong evidence of the non-monotone effect of uncertainty predicted by the theory. Instrumental variable estimates and several robustness tests confirm this result.

How does our paper relate to previous explanations of the Asian crisis? While our focus on the role of uncertainty is unique, our empirical results are broadly consistent with the prevailing idea of the Asian crisis as rooted in the financial conditions of the corporate and banking sectors. Indeed, this is probably the reason why GDP growth forecasts are the best measure of exchange rate fundamentals in our sample, proving to be empirically superior to forecasts of other macroeconomic variables (such as inflation or the current account) that have played a key role in other crisis episodes. The statistically significant coefficient of the interaction between U.S. interest rates and short-term external debt, which we also find, is consistent with the role of internationally illiquid banks emphasized by Chang and Velasco (2001) as well as with that of foreign currency borrowing by domestic firms in the model of Aghion, Bacchetta, and Banerjee (2004). Our paper, however, cannot be seen as a direct empirical test of these models, which are characterized by multiple equilibria. Developing a framework for testing the role of uncertainty in models with multiple equilibria is a topic for future research, possibly using regime switching econometric techniques as in Jeanne and Masson (2000).

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. In addition, it might be worthwhile exploring whether these variables can enhance the predictive power of early warning systems, which are often based on past fundamentals. Indeed, drawing on a previous version of our paper, Bannier (2006) applies our empirical framework to Mexican data obtaining qualitatively similar results and Köhler (2007) uses it to explore the role of uncertainty about fundamentals in sudden stops.

Appendix

A A simple first-generation model

In this section we show that the basic equations of first-generation models can be collapsed into equation (2). First-generation models are based on the following core set of assumptions:

$$
\begin{aligned}
m_t - p_t &= a + b y_t - c i_t && (4)\\
p_t - p_t^* &= e_t && (5)\\
E_t(e_{t+1}) &= e_t + i_t - i_t^* && (6)\\
m_t &= \ln(R_t + D_t) && (7)
\end{aligned}
$$

where m_t, y_t and p_t represent the logarithms of domestic money stock, income, and prices; a, b, and c are positive constants; e_t is the logarithm of the nominal exchange rate, defined as units of domestic currency for one unit of foreign currency; i_t is the domestic nominal interest rate; E_t is the expectation operator conditional on the information at time t; D_t and R_t denote the stocks of domestic credit and international reserves; and an asterisk identifies foreign variables. Equations (4), (5) and (6) respectively represent the money market equilibrium condition, the purchasing power parity, and the uncovered interest parity.

Equations (4)-(6) can be rearranged into the standard forward-looking exchange rate equation:

$$(1+c)\,e_t - cE_t\,(e_{t+1}) = m_t + v_t \ , \tag{8}$$

where $v_t \equiv ci_t^* - a - by_t - p_t^*$ denotes an "all-inclusive" velocity term. We also write $f_t \equiv m_t + v_t$, where f_t gathers the fundamentals that determine the behavior of the exchange rate. Clearly, an increase in f_t would correspond to a deterioration in the fundamentals: in other words, a larger f_t causes a depreciation of the equilibrium exchange rate.

If the exchange rate were free to fluctuate, we could determine its equilibrium value by solving the difference equation (8), thereby obtaining:

$$e_t = \frac{1}{1+c}\sum_{j=0}^{+\infty}\left(\frac{c}{1+c}\right)^j E_t\,(f_{t+j}) \ . \tag{9}$$

We also assume that public authorities keep the exchange rate fixed at a given peg $\overline{e}$ until their international reserves are larger than some lower bound $\underline{R} \in \mathbb{R}$, while at $R_t = \underline{R}$ they abandon the peg leaving the exchange rate free to fluctuate.

Define the logarithm of money stock after a devaluation as $\underline{m_t} = \ln(\underline{R} + D_t)$. Analogously, denote with $\underline{f_t}$ the fundamentals after a devaluation (i.e. $\underline{f_t} = \underline{m_t} + v_t$). The equilibrium exchange rate conditional on a successful attack (the so-called *shadow exchange rate*), can be obtained by plugging $\underline{f_t}$ into equation (9); namely:

$$\widetilde{e}_t = \frac{1}{1+c}\sum_{j=0}^{+\infty}\left(\frac{c}{1+c}\right)^j E_t\left(\underline{f_{t+j}}\right) \ . \tag{10}$$

Let us assume that fundamentals follow the random process: $\underline{f_t} = y + \underline{f_{t-1}} + \lambda_t$, with $y > 0$ and with λ_t being the realization of the random variable: $\Lambda_t \sim Norm\,(0, 1/\alpha)$. This stochastic process implies that fundamentals deteriorate over time, as a result of, e.g., excess growth in the velocity term or in the domestic credit. In order to keep the peg fixed at the value $\overline{e}$ while f_t grows, public authorities have to adjust R_t so that the money market equilibrium condition (4) continues to hold. Taking expectations, we obtain $E_t\left(\underline{f_{t+j}}\right) = jy + \underline{f_t}$ and, plugging this average into equation (10), we find: $\widetilde{e}_t = cy + \underline{f_t}$.

The probability of a speculative attack, $\Pr(\widetilde{e}_{t+1} \geq \overline{e})$, becomes: $\Pr(\Lambda_{t+1} \geq l_t)$, where $l_t = \overline{e} - y\,(1-c) - \underline{f_t}$. Note that, to maintain notation consistency with Section 2.1, it is enough to define define $\theta_t = y - \lambda_t$, so that its distribution is $\Theta_t \sim Norm(y, 1/\alpha)$ and where a larger θ_t means better fundamentals. With this notation change, the time-t probability of a speculative attack is $\Pr(\Theta_t \leq k_t)$ (i.e. equation (2)), where $k_t = y - l_t$.

B Thresholds values of Proposition 2

Morris and Shin (2004) and Hellwig (2002) show that the trigger points θ^* and x^* solve:

$$\begin{aligned} 0 &= \theta^* - \Phi\left[\sqrt{\beta}\,(x^* - \theta^*)\right] \\ 0 &= D \cdot \Phi\left[\sqrt{\alpha+\beta}\left(\theta^* - \frac{\alpha}{\alpha+\beta}y - \frac{\beta}{\alpha+\beta}x^*\right)\right] - t\,. \end{aligned} \tag{11}$$

If $\beta \geq \frac{\alpha^2}{2\pi}$, the system (11) yields a unique solution (even though it is not in a closed form). The share of speculators attacking the currency is: $\Pr(X_i \leq x^* \mid \theta) = \Phi\left[\sqrt{\beta}\,(x^* - \theta)\right]$; hence:

$$\frac{d\Pr(X_i \leq x^* \mid \theta)}{d\alpha} = \sqrt{\beta} \cdot \phi\left[\sqrt{\beta}\,(x^* - \theta)\right] \cdot \frac{dx^*}{d\alpha}\,, \tag{12}$$

where ϕ is the pdf of the standard normal distribution. Therefore, the share of attackers and the equilibrium trigger point x^* have the same threshold. By differentiating the equations in the system (11), it is easy to find that $\frac{dx^*}{d\alpha} \leq 0 \Leftrightarrow y > s_1$, where

$$s_1 = \frac{\alpha+\beta}{\alpha+2\beta}\theta^* + \frac{\beta}{\alpha+2\beta}x^*\,,$$

which shows that the threshold s_1 is a weighted average of the equilibrium trigger points x^* and θ^* (see Prati and Sbracia, 2002, for details).

Computing the derivative of the share of attackers with respect to β is somewhat more cumbersome. Clearly:

$$\frac{d\Pr(X_i \leq x^* \mid \theta)}{d\beta} = \phi\left[\sqrt{\beta}\,(x^* - \theta)\right] \cdot \left(\sqrt{\beta}\frac{dx^*}{d\beta} + \frac{x^* - \theta}{2\sqrt{\beta}}\right)\,. \tag{13}$$

Thus, the share of attackers and the equilibrium trigger point x^* now have different thresholds. In particular, the share of attackers is decreasing in β if $\theta > s_2$, where

$$s_2 = 2\beta\frac{dx^*}{d\beta} + x^*\,,$$

with:

$$\frac{dx^*}{d\beta} = \frac{\alpha\sqrt{\beta} + \alpha\beta\phi_0}{2\beta\,(\alpha+\beta)\,\alpha\beta\phi_0}y - \frac{2\alpha\sqrt{\beta} + \beta\sqrt{\beta} - \alpha^2\phi_0}{2\beta\,(\alpha+\beta)\,\alpha\beta\phi_0}x^* + \frac{\theta^*}{2\beta}\,,$$

and with ϕ_0 denoting $\phi\left[\sqrt{\beta}\,(x^* - \theta)\right]$. Analogously, the equilibrium trigger point is increasing in β if $y > s_3$, where

$$s_3 = \frac{A+C}{A+B}x^* + \frac{A-C}{A+B}\theta^*\,,$$

with $A = \alpha\sqrt{\beta}$, $B = \alpha\beta\phi_0$, and $C = \alpha\sqrt{\beta} + \beta\sqrt{\beta} - \alpha^2\phi_0$; thus, s_3 is a linear combination of x^* and θ^* (again, Prati and Sbracia, 2002, contains the algebraic details).

C Interpreting the forecasts

In this section, we show that, under general assumptions on how forecasters make their predictions, the mean forecast measures expected fundamentals and the variance of the forecasts is increasing in the variance of public and private information.

Consistently with Section 2, the variable $\Theta \sim Norm(y, 1/\alpha)$ describes fundamentals, and each agent i receives a private signal $x_i = \theta + \varepsilon_i$, with $\varepsilon_i \sim Norm(0, 1/\beta)$ (ε_i and ε_j independent given θ for $i \neq j$). Bayesian updating implies:

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

Let us assume that each agent i reports to Consensus Economics the mean of his ex-post beliefs (*honest forecasting*): $f_{h,i}(x_i) = (\alpha y + \beta x_i) / (\alpha + \beta)$. The mean of the individual forecasts of n forecasters then is:

$$f_h(x_1, ..., x_n) = \frac{\sum f_{h,i}(x_i)}{n} = \frac{\alpha}{\alpha + \beta} y + \frac{\beta}{\alpha + \beta} \frac{\sum x_i}{n} . \tag{14}$$

Given θ, for n that goes to $+\infty$ this random variable converges to:

$$\mu_h(\theta) = E[f^e(X_1, ..., X_n) \mid \theta] = \frac{\alpha}{\alpha + \beta} y + \frac{\beta}{\alpha + \beta} \theta . \tag{15}$$

If n is sufficiently large, by using the mean forecast in the empirical analysis we use a variable that is influenced by θ and y. Since $E(\Theta) = y$, then $E[\mu_h(\Theta)] = y$. Thus, *on average* the mean of individual forecasts is identically equal to the public signal y and does not depend 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 in any systematic way on α and β.

Equations (14) and (15) show that the mean forecast is an unbiased but imprecise measure of y. It is affected by a sample error, due to the fact that we observe only a finite number of forecasters (as apparent from equation (14)). Even absent the sample error (i.e. for $n \to +\infty$), however, it would still be affected by a forecast error, as shown by the fact that equation (15) can be rewritten as $\mu_h(\theta) = y + \varepsilon$, where $\varepsilon = (\theta - y) \beta / (\alpha + \beta)$. These considerations, then, corroborate the importance of using IV estimates in our analysis to address this measurement problem.

Let us now turn to the variance of the individual forecasts:

$$s_h^2(x_1, ..., x_n) = \frac{\beta^2}{(\alpha + \beta)^2} \frac{\sum (x_i - \overline{x})^2}{n} , \tag{16}$$

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

$$\sigma_h^2(\theta) = E\left[\sigma_h^2(X_1, ..., X_n) \mid \theta\right] = \frac{\beta}{(\alpha + \beta)^2} . \tag{17}$$

For n sufficiently large, a change in y affects the mean of the individual forecasts μ_h but does not affect their variance σ_h^2, which only depends on α and β.

Besides the measurement problem due to the sample error shown by equation (16), it is apparent from equation (17) that while an increase in α always implies a decrease in σ_h^2, an increase in β does not necessarily reduce σ_h^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. On the other hand, for given messages x_i, the rise in β increases the weight of the private messages in the individual predictions, making them more heterogeneous among the forecasters. The first (second) effect dominates when $\beta > \alpha$ ($\beta < \alpha$).

Honest forecasting corresponds to the *forecasters' optimal strategy* when their payoffs depend solely and symmetrically on the accuracy of their predictions (i.e., when the payoff is of the type: $a \cdot |\theta - f_{h,i}(x_i)|^b$, for $a < 0$ and $b > 0$). Ottaviani and Sørensen (2006) consider two other possible strategic objectives of forecasters: a *winner-take-all contest*, in which the forecaster whose prediction turns out to be closest to the fundamental wins a prize, and a *reputational-cheap-talk game*, an incomplete information game in which forecasters want to signal that they are the best informed type. These two games share the following properties: (i) on average, the mean forecasts is equal to the expected fundamental y (as for honest forecasting); (ii) the variance of the forecasts is *always* a monotonic function of the variances of public and private information. Note that a key result from both these strategies is that *property (ii) holds for any* α *and* β; in particular, it holds also in the case $\beta < \alpha$.[24]

It is not necessary to place any restriction on α and β to grant that property (ii) holds also for a number of other possible forecasting strategies, such as *maximum likelihood forecasting* or *if information is only public.* For the latter case, which is consistent with the first-generation model of Section 2.1, Laster, Bennet, and Geoum (1999) show that agents' equilibrium forecasts always differ (even if their information set is the same) and their distribution matches the distribution of Θ.

Thus, besides the case of honest forecasting, the condition $\beta > \alpha$ is not necessary for the variance of the forecasts to be increasing in the variance of public and private information. Nonetheless, we can attempt an estimate of the relative values of α and β to check whether $\beta > \alpha$. Specifically, we can regress the mean forecasts of GDP growth on actual GDP growth and a constant. By equation (15), the slope of this regression, which we denote by $\hat{\rho}_1$, provides an estimate of $\beta/(\alpha+\beta)$. This is just a rough estimate of the order of magnitude

[24] If agents use a *convex forecast strategy* of the kind: $f_{p,i}(x_i) = [1 - P(\alpha,\beta)] \cdot y + P(\alpha,\beta) \cdot x_i$, with $P(\alpha,\beta) \in (0,1)$, it is immediate to check that property (i) always holds. For the variance of the forecasts to be increasing in the variances of public and private information, instead, some restrictions on the weight $P(\alpha,\beta)$ are required. Specifically, one needs: (a) $\partial_\alpha P(\alpha,\beta) < 0$: i.e. if public information are more precise, the weight assigned to the public signal must increase; and (ii) $\partial_\beta P(\alpha,\beta) < P(\alpha,\beta)\beta^{-1}$: i.e. if private information are more precise, the weight placed on the private signal can increase (the condition is consistent with $\partial_\beta P(\alpha,\beta) > 0$), but "not too quickly."

of $\beta/(\alpha+\beta)$, because this regression implicitly assumes that this ratio is constant over time. Interestingly, we find that the value of $\hat{\rho}_1$ depends on the forecast horizon, and increases as the forecast horizon shortens. With the 12-month forecast horizon used in this paper, we reject the hypothesis $\hat{\rho}_1 < 0.5$ (i.e. $\beta < \alpha$) for 4 out of 6 countries, while for Hong Kong and Singapore we obtain values of $\hat{\rho}_1$ in the order of 0.2. These results are confirmed when we replace the mean with the median forecast as the dependent variable.[25]

D Additional robustness checks

Additional forecast and actual data. In the specification search, we did not limit our analysis to forecasts of GDP growth. We experimented with other forecasts of macroeconomic variables from Consensus Economics — inflation, current account balance, trade balance, and exports — but these variables did not perform as well as GDP growth and, when we included measures of the mean and variance of expected GDP growth in the regression, hardly any other forecast variable was significant. This is not surprising once we recall that imbalances in the financial and corporate sectors are believed to be at the root of the Asian crisis and note that the economic conditions of these sectors are more likely to be reflected in GDP growth forecasts than in forecasts of other macroeconomic variables.

We also considered other regressors from actual, rather than forecast, data. A non-exhaustive list we experimented with includes: GDP growth, inflation, international reserves, and the ratio of M2 to international reserves. However, none of these variables had a significant effect on exchange rate pressures once we included the mean and variance of expected GDP growth in the regression. In particular, when we tried actual GDP growth — whose inclusion in the empirical model would have potentially allowed to attenuate the measurement problem, since equation (15) shows that the mismeasurement depends on actual fundamentals — this variable was not significant while $\hat{\gamma}_1$ and $\hat{\gamma}_2$ remained significant and with the predicted sign.

Robustness to dynamic specification and endogenous regressors. As it is common in the literature, when we test the global-game model 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 have corrected for serial correlation of the errors in our regressions by including a country-specific AR(1) term rather than estimating a truly dynamic specification. In a possible dynamic extension of the theory, 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 global-game model is beyond the scope of this paper, we can estimate a dy-

[25] We thank a referee both for pointing out the measurement problem affecting the mean forecast and for suggesting this regression.

namic version of equation (3) 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 2 confirming, in particular, 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 baseline regression and are both statistically significant at the 1 percent confidence level.

The GMM estimator provides also an alternative way of checking for the robustness of our results to the potential endogeneity of the one-period-lagged mean and variance of the forecasts by instrumenting them 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 baseline 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 $\hat{\gamma}_1$ 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 with the precision of public and private information. We allow for this possibility by estimating $\hat{\gamma}_1$ recursively with state-space techniques. We find 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 finding that the recursive estimate of $\hat{\gamma}_2$ changes over time but remains always positive and significantly different from zero.[26]

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,

[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 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.

reflecting changes in the parameters in k_t, s_1, and s_2 in the models of Section 2 or, for instance, because investors might have simply 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 (3) recursively. 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, allowing for time-varying thresholds has little effect on $\hat{\gamma}_1$ and $\hat{\gamma}_2$, which remain significant and with the predicted sign. 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 ($\underline{\hat{\gamma}_j}_{GDP,t-1}$). We find 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.

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Table 1. Values of the index of exchange rate pressure for different levels of the average and variance of GDP growth forecasts [1]

	"Bad" average forecast	"Good" average forecast
High forecast variance *(obs)*	-0.157 *(259)*	0.797 *(113)*
Low forecast variance (obs)	0.348 *(113)*	-0.163 *(259)*

[1] The average forecasts is "bad" ("good") if it is below (above) the country's median GDP growth forecast; the forecasts variance is high if it is above (below) the median sample value. The index of exchange rate pressure is the sum of the normalized values of: (*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 (see Section 3.1).

Table 2. Exchange Rate Pressure (**IND3** Index) Estimates
(SUR estimates; standard errors in parenthesis; sample: 1995:03 - 2005:04)[1]

	Thailand	*Indonesia*	*South Korea*	*Malaysia*	*Singapore*	*Hong Kong*	*Panel*[2]
$\gamma_{0,j}$	-3.130	-2.232 *	-12.807 ***	-8.517 ***	-10.446 ***	4.962 *	Fixed effects:
	(2.477)	(1.238)	(3.171)	(2.356)	(3.212)	(2.571)	YES
γ_1	-0.361 ***						-1.426 ***
	(0.082)						(0.313)
γ_2	0.633 ***						2.803 ***
	(0.067)						(0.890)
γ_3	0.025						-0.323
	(0.064)						(0.331)
γ_4 *1000	0.441 ***						1.035 ***
	(0.104)						(0.345)
$\underline{\gamma}_j$	2.508 ***	0.469	1.661	4.078 ***	1.024	-0.947	7.310 ***
	(0.746)	(0.777)	(1.129)	(1.044)	(0.957)	(1.223)	(0.350)
$\gamma_{5,j}$	0.057 *	0.041 ***	0.163 ***	0.114 ***	0.111 ***	-0.044 *	-0.029
	(0.029)	(0.015)	(0.036)	(0.025)	(0.032)	(0.025)	(0.032)
ρ_j	0.334 ***	0.381 ***	0.500 ***	0.460 ***	0.203 ***	0.189 **	0.248 ***
	(0.074)	(0.088)	(0.079)	(0.069)	(0.072)	(0.083)	(0.071)
R^2	0.357	0.698	0.557	0.424	0.108	0.377	0.329
DW	1.678	2.142	1.568	1.872	1.784	2.157	2.050
Observations	123	123	123	123	123	123	180

[1] Data are monthly. Three (***), two (**) and one (*) stars mark statistical significance respectively at one, five, and ten percent levels. [2]The panel sample is 1995:03 - 1997:07. Panel Durbin-Watson statistics are calculated using the within residuals as suggested by Bhargava et al. (1982).

Standard Errors for $\underline{\gamma}_j$ are calculated using the delta method.

Table 3. Exchange Rate Pressure (**IND3** Index) Estimates
(3SLS estimates using United States consumption instruments; standard errors in parenthesis; sample: 1995:03 - 2005:04)[1]

	Thailand	*Indonesia*	*South Korea*	*Malaysia*	*Singapore*	*Hong Kong*	*Panel*[2]
$\gamma_{0,j}$	-2.639	0.280	-22.682 ***	-5.344 *	-5.422	7.712 ***	Fixed Effects:
	(3.179)	(1.563)	(5.790)	(2.751)	(3.600)	(2.687)	YES
γ_1	-0.472 ***						-1.928 ***
	(0.125)						(0.488)
γ_2	0.788 ***						4.537 ***
	(0.105)						(1.327)
γ_3	0.128 *						-0.298
	(0.068)						(0.393)
γ_4 *1000	0.345 ***						0.832 *
	(0.105)						(0.431)
$\underline{\gamma}_j$	1.785 *	2.006 *	2.765	4.329 ***	0.906	-2.477 *	7.013 ***
	(0.960)	(1.130)	(1.925)	(1.301)	(1.131)	(1.371)	(0.328)
$\gamma_{5,j}$	0.048	0.017	0.285 ***	0.082 ***	0.060 *	-0.079 ***	-0.049
	(0.037)	(0.017)	(0.063)	(0.028)	(0.036)	(0.026)	(0.039)
ρ_j	0.444 ***	0.356 ***	0.833 ***	0.496 ***	0.265 ***	0.099	0.351 ***
	(0.087)	(0.094)	(0.065)	(0.080)	(0.086)	(0.089)	(0.081)
R^2	0.366	0.683	0.584	0.433	0.112	0.378	0.332
DW	1.898	2.154	1.939	2.039	1.922	1.990	1.969
Observations	123	123	123	123	123	123	180

[1] Data are monthly. Three (***), two (**) and one (*) stars mark statistical significance respectively at one, five, and ten percent levels. [2]The panel sample is 1995:03 - 1997:07. Panel Durbin-Watson statistics are calculated using the within residuals as suggested by Bhargava et al. (1982).

Standard Errors for $\underline{\gamma}_j$ are calculated using the delta method.

Table 4. Exchange Rate Pressure (**IND3** Index) Estimates
(3SLS estimates using United States unemployment instruments; standard errors in parenthesis; sample: 1995:03 - 2005:04)[1]

	Thailand	*Indonesia*	*South Korea*	*Malaysia*	*Singapore*	*Hong Kong*	*Panel*[2]
$\gamma_{0,j}$	-2.763	-0.855	-24.861 ***	-5.425 *	-5.021	6.918 **	Fixed effects:
	(3.222)	(1.463)	(5.729)	(2.793)	(3.632)	(2.776)	YES
γ_1				-0.384 ***			-2.254 ***
				(0.125)			(0.457)
γ_2				0.738 ***			5.487 ***
				(0.105)			(1.301)
γ_3				0.096 ***			-0.245
				(0.068)			(0.411)
γ_4 *1000				0.332 ***			0.658
				(0.109)			(0.450)
$\underline{\gamma}_j$	1.294	1.031	2.021	4.034 ***	0.043	-3.360 **	6.943 ***
	(1.067)	(1.111)	(2.144)	(1.453)	(1.390)	(1.613)	(0.246)
$\gamma_{5,j}$	0.044	0.023	0.304 ***	0.078 ***	0.050	-0.075 ***	-0.063
	(0.038)	(0.016)	(0.062)	(0.029)	(0.036)	(0.027)	(0.041)
ρ_j	0.447 ***	0.336 ***	0.850 ***	0.500 ***	0.261 ***	0.117	0.370 ***
	(0.088)	(0.093)	(0.060)	(0.080)	(0.086)	(0.089)	(0.081)
R^2	0.367	0.699	0.588	0.440	0.101	0.364	0.329
DW	1.913	2.145	1.932	2.067	1.923	2.003	2.006
Observations	123	123	123	123	123	123	180

[1] Data are monthly. Three (***), two (**) and one (*) stars mark statistical significance respectively at one, five, and ten percent levels. [2]The panel sample is 1995:03 - 1997:07. Panel Durbin-Watson statistics are calculated using the within residuals as suggested by Bhargava et al. (1982).

Standard Errors for $\underline{\gamma}_j$ are calculated using the delta method.

Table 5. Exchange Rate Pressure (**IND3** Index) Estimates
(3SLS estimates using G6-composite consumption instruments; standard errors in parenthesis; sample: 1995:03 - 2005:04)[1]

	Thailand	Indonesia	South Korea	Malaysia	Singapore	Hong Kong	Panel[2]
$\gamma_{0,j}$	-2.954 (3.211)	0.040 (1.558)	-22.864 *** (5.769)	-5.503 ** (2.771)	-5.101 (3.614)	7.709 *** (2.689)	Fixed effects: YES
γ_1	-0.484 *** (0.125)						-1.895 *** (0.487)
γ_2	0.801 *** (0.106)						4.433 *** (1.324)
γ_3	0.126 * (0.068)						-0.302 (0.395)
γ_4 *1000	0.3384 *** (0.105)						0.827 * (0.435)
$\underline{\gamma}_j$	1.724 (0.943)	1.838 (1.096)	2.656 (1.882)	4.198 *** (1.282)	0.941 (1.111)	-2.381 * (1.348)	6.952 *** (0.337)
$\gamma_{5,j}$	0.051 (0.038)	0.020 (0.017)	0.287 *** (0.063)	0.084 *** (0.028)	0.057 (0.036)	-0.078 *** (0.026)	-0.050 (0.039)
ρ_j	0.448 *** (0.087)	0.356 *** (0.093)	0.832 *** (0.065)	0.497 *** (0.080)	0.267 *** (0.086)	0.100 (0.089)	0.357 *** (0.082)
R^2	0.366	0.684	0.585	0.433	0.112	0.378	0.282
DW	1.903	2.149	1.938	2.037	1.926	1.993	1.985
Observations	123	123	123	123	123	123	180

[1] Data are monthly. Three (***), two (**) and one (*) stars mark statistical significance respectively at one, five, and ten percent levels. [2]The panel sample is 1995:03 - 1997:07. Panel Durbin-Watson statistics are calculated using the within residuals as suggested by Bhargava et al. (1982).

Standard Errors for $\underline{\gamma}_j$ are calculated using the delta method.

Table 6. Exchange Rate Pressure (**IND3** Index) Estimates
(3SLS estimates using G6-composite unemployment instruments; standard errors in parenthesis; sample: 1995:03 - 2005:04)[1]

	Thailand	*Indonesia*	*South Korea*	*Malaysia*	*Singapore*	*Hong Kong*	*Panel*[2]
$\gamma_{0,j}$	-2.632	-0.550	-23.834 ***	-5.419 **	-5.125	7.165 **	Fixed effects:
	(3.174)	(1.428)	(5.830)	(2.703)	(3.609)	(2.786)	YES
γ_1				-0.372 ***			-2.107 ***
				(0.123)			(0.446)
γ_2				0.741 ***			4.998 ***
				(0.097)			(1.261)
γ_3				0.102			-0.265
				(0.068)			(0.386)
γ_4 *1000				0.315 ***			0.821 *
				(0.110)			(0.421)
$\underline{\gamma}_j$	1.260	1.150	1.739	4.109 ***	0.207	-3.726 **	7.040 ***
	(1.060)	(1.082)	(2.190)	(1.415)	(1.367)	(1.648)	(0.268)
$\gamma_{5,j}$	0.041	0.019	0.289 ***	0.077 **	0.051	-0.079 ***	-0.051
	(0.037)	(0.016)	(0.063)	(0.028)	(0.036)	(0.027)	(0.039)
ρ_j	0.438 ***	0.333 ***	0.830 ***	0.482 ***	0.251 ***	0.122	0.330 ***
	(0.088)	(0.093)	(0.065)	(0.081)	(0.086)	(0.090)	(0.080)
R^2	0.367	0.696	0.587	0.442	0.103	0.356	0.323
DW	1.902	2.145	1.926	2.036	1.906	2.000	1.926
Observations	123	123	123	123	123	123	180

[1] Data are monthly. Three (***), two (**) and one (*) stars mark statistical significance respectively at one, five, and ten percent levels. [2]The panel sample is 1995:03 - 1997:07. Panel Durbin-Watson statistics are calculated using the within residuals as suggested by Bhargava et al. (1982).

Standard Errors for $\underline{\gamma}_j$ are calculated using the delta method.