IMF Programs and Economic Crisis:
An Empirical Study of Transition

Patrick Conway
Department of Economics
Gardner Hall, CB 3305
University of North Carolina
Chapel Hill, NC 27599-3305
patrick_conway@unc.edu

Third revision: 5 January 2000

Abstract.

According to the IMF Articles of Agreement, IMF resources are disbursed in response to balance of payments problems and should lead to a solution to those problems. While a number of authors have assessed the impact of IMF programs on performance (see, for example, Khan 1990, Conway 1994, Bird 1995, Killick 1995, Fischer 1997), there has been little attention to the dynamic of participation in IMF programs. Is this transition forecast in the Articles of Agreement in fact observed?

In this paper I address this question for a sample of 90 developing countries over the period from the beginning of 1974 to the end of 1992. Simple estimates of the impact of IMF programs on crisis yield negative conclusions about those programs. Estimates that control for country- and time-specific variation in the data indicate that the transition forecast in the Articles of Agreement does exist. However, there is a reduction of its effectiveness as countries spend more and more time in IMF programs.

Key words: IMF programs, economic crisis, conditionality
JEL classifications: F32, F34, O19

Thanks to Nadeem ul Haq for stimulating my interest in this question, to Stanley Black and Alfred Field for their comments, and to seminar participants at Cornell University, Duke University and the University of North Carolina at Chapel Hill for perceptive critiques.

The debate over the role of the International Monetary Fund (IMF) in the recent economic crises in East Asia has placed a spotlight on an ongoing debate among academics and developing-country policy-makers about the impact of IMF involvement in developing-country policy formulation. Developing countries turn to the IMF for financial assistance in times of foreign-exchange shortfalls. While they receive this assistance, in accordance with the IMF Articles of Agreement, they are often called upon as a precondition of this assistance to undertake economic reforms proposed by IMF staff. These reforms are in principle designed to speed the developing country's departure from its shortfall position. Critics have recently claimed, however, that these reforms have in fact worsened the plight of the borrowing country and brought nearer the date of the next shortfall.[1]

The impact of participation in an IMF program on the developing country's economic performance has been studied previously in the empirical literature (see, for example, Khan 1990, Conway 1994, Bird 1995, Killick 1995, and Fischer 1997). I'm naturally partial to the results in Conway (1994), with its conclusions of initial worsening of real economic growth, improvement in the current account and initial worsening of the investment/GDP ratio, but other authors will dispute these conclusions. The Conway (1994) analysis was also unique in that it allowed for a rudimentary dynamic in the impact of IMF program participation – the contemporaneous effect was allowed to differ from the lagged effect, and the estimation results demonstrated a significant

[1] For critics of the IMF's advice in the East Asian crisis, see Sachs (1997), the Economist (1997), Feldstein (1998), and Rosett (1999). For similar criticism on its handling of the Russian crisis, see Sachs (1998). These criticisms are in the earlier spirit of Killick (1984, 1995) that concluded IMF programs may well worsen a crisis by not properly targeting and remedying its sources.

difference. In both economic growth and investment behavior, for example, a significant positive lagged effect of the IMF program worked to counter the contemporaneous negative effect.

In this paper I expand upon the dynamic inherent in the use of IMF resources. They are disbursed in response to external imbalances in the developing country, and the joint provision of financing and adoption of IMF preconditions is designed to lead to an end to those imbalances. Adopting the present policy terminology, economic crisis leads to participation in IMF programs, and this participation should lead to an end to this crisis. Is the dynamic transition from crisis to non-crisis through participation in IMF programs in fact observed?

There are two substantive departures from previous work on IMF programs employed in this paper. First, I use a number of empirical indicators of "economic crisis" to provide a measurement of the effect of participation in IMF programs. The resulting estimates indicate the change in the conditional probability of ending economic crisis by participating. Second, I consider as well the impact of IMF participation on the probability that crisis will recur. These substantive departures require methodological departures, as well, with use of survival analysis and estimation of country-specific "frailties" more commonly found in biostatistical analysis.

I address this question for a sample of 90 developing countries on a quarterly basis over the period from the beginning of 1974 to the end of 1992. The conclusions from this statistical analysis yield a mixed verdict on participation in IMF programs. On the positive side, participation in an IMF program while in crisis leads to a significant improvement in the probability that the country will end the crisis – or alternatively put, the country can expect to see an end to its crisis at an earlier date. On the negative side, there is significant evidence that the greater the time spent in preceding years in IMF programs the lower is the probability of ending

the crisis in any period; in other words, there is evidently a weakening of a country participating in repeated or prolonged IMF programs that exists independently of other country-specific effects. The empirical results finally leave us with a mystery, for none of the indicators of economic crisis considered is especially strongly associated with participation in IMF programs.

I. Framing the question.

I begin from a maintained premise: that measurement of the impact on developing countries of participating in IMF programs cannot be separated from the motivation of the developing country to participate in that program.

$$y_{jt} = \mu_{yj} + x_{j\{t\}}b + I_{j\{t\}}g + e_{yjt} \quad (1)$$

$$I_{it} = \mu_{Ij} + Z_{j\{t\}}d + Y_{j\{t-1\}}f + I_{j\{t-1\}}h + e_{Ijt} \quad (2)$$

Conway (1994) provides an initial pass at modeling this interlinkage. Each performance variable y_{jt} for country j in period t is modeled as determined by a country-specific term μ_{yj}, a number of exogenous variables, both contemporaneous and lagged, denoted by the matrix $x_{j\{t\}}$, and the participation in IMF programs in contemporaneous and previous periods $I_{j\{t\}}$.[2] Participation in IMF programs is modeled through a reduced-form specification of the determinants of participation. Both "demand-side" factors associated with the participating country's economic

[2] The b and g vectors are sized to conform with the dimensions of the $x_{j\{t\}}$ and $I_{j\{t\}}$ matrices. The notation {t} indicates that the variable will be represented either by its current value or lagged values in the explanatory equation – for example, $x_{j\{t\}}$ could include as columns the variables x_{jt}, x_{jt-1}, x_{jt-2}, and so on. Each will have its own element in the vector b as coefficient.

situation (represented by various performance variables y_{jt-1} collected into the matrix $Y_{j\{t-1\}}$) and "supply-side" factors of importance to IMF decision-makers (included in $Z_{j\{t\}}$) were found to be significant determinants of participation, as was participation in the previous periods. ($I_{j\{t-1\}}$) . The specification was limited in that $I_{j\{t\}}$ was the only channel for IMF participation to affect the economic performance: notably, the coefficients on the exogenous variables $x_{j\{t\}}$ were held equal in participation and non-participation periods. In equation (2) the terms $Y_{j\{t-1\}}$f capture in hedonic form the impact of economic crisis on IMF participation. While the coefficients of f by and large take the correct sign for such an interpretation, there is an element of circularity in defining economic crisis in this way.

The modeling strategy in this paper builds upon and amends that outlined above. First, a number of indicators of economic crisis are derived independently of IMF participation. Second, the conditional probabilities of entry into and exit from crisis are examined in a switching-regression structure (Maddala (1983), ch. 8) with the impact of IMF participation on these probabilities derived jointly with changes in the policy and market structure due to the crisis – or non-crisis. Third, the conditional probabilities of entry into and exit from crisis are calculated through an application of survival analysis; these have a natural dynamic application in terms of the length in periods of episodes of crisis – or non-crisis – and the impact of IMF participation on the length of those episodes.

II. **Defining and measuring economic crisis and IMF program participation**.

There is no agreed-upon definition of external crisis in the economics literature. The stock of foreign-exchange reserves scaled by the country's demands for them is one measure suggested in Agenor and Montiel (1996, ch. 6) and Obstfeld and Rogoff (1996, ch. 8). Corsetti, Pesenti and Roubini (1998a) and Kaminsky and Reinhart (1999) create indices based upon a weighted average of (the negative of) foreign exchange reserve growth and depreciation of the nominal exchange rate, with outlying positive values of the index interpreted as the onset of crisis. These indices have two disadvantages from the standpoint of this study: first, they pinpoint the beginning of a crisis period but not necessarily an ending point; second, they include exchange rate depreciation which can as easily be thought of as a response to a crisis as the crisis itself.

In this paper I report in detail the results obtained from three definitions of crisis. The first (C_{Rit}) is based upon observed ratios of foreign exchange reserves to imports below a critical value.[3] The second (C_{Git}) is based upon observed growth rates in foreign exchange reserves that fall below a critical value. The third (C_{Iit}) is based upon observed inflation rates that exceed a critical value.

In the period from the beginning of 1973 to the end of 1992, the IMF offered five types of program agreements.[4] The "first tranche" agreement was a decision by the borrowing country to

---

[3] An initial correction was performed to these ratios to account for the reduced reserve-holding needs of the CFA franc zone and of Liberia. These African nations have mean reserve/import ratio .55 less than the other countries in the sample. In calculating a measure of crisis, I adjusted the observed ratios for these countries upward in each period by that amount.

[4] In some cases, countries were able to negotiate a hybrid of more than one of these types. There were in addition other facilities offered by the International Monetary Fund as well that provided financing without the notion of conditionality associated with the programs cited. These other facilities included the Compensatory Financing Facility, the Buffer Stock Financing Facility and the Oil Facility, and participation in these is not considered here. The Supplementary Financing Facility, established in 1979,

draw down its hard-currency position at the IMF. The "stand-by" was an agreement by the IMF to allow the member country to draw down a percentage of its quota, also in hard currency. Repayment (or, more properly, redeposit) under these stand-by agreements occurred usually over a single year, although exceptions were made in some circumstances. The "extended fund facility", or EFF, was an agreement to allow members to draw down some percentage of their quota, but with repayment scheduled over a longer (often 3-year) period. These agreements were first extended in September 1974. The "structural adjustment facility" (SAF) and "extended structural adjustment facility" (ESAF) were introduced in 1986. The SAF had a similar repayment period to the EFF, but was extended to members whose balance of payments difficulties were viewed to be more systemic, or structural, in origin. The ESAF was designed to provide longer repayment periods than the SAF for countries with especially difficult structural adjustments to complete. Participation in any one of these during the period is considered participation in an IMF program for the purposes of this study.[5]

Participation in some type of IMF agreement was characteristic of 34.1 percent of the quarterly observations, as Table 1 indicates. By far the most common of these was the stand-by arrangement (17.6 percent of the total sample), with EFF, first tranche, SAF and ESAF programs observed less frequently. Hybrid programs (for example, a stand-by/SAF combination) were observed in 1.7 percent of the total sample.

Given the frequency of IMF participation I define the crisis variables so that the critical

---

provided added resources (from borrowed funds rather than subscriptions) to participating countries; this facility was only available in tandem with one of the five programs noted in the text.

[5] A persuasive argument can be made that first tranche agreements are significantly different from the others in terms of conditionality and relation to crisis. I have redone the analysis here with a data set excluding those programs and obtain very similar results. These are available on demand.

value of the indicator places 34.1 percent of the sample in the category of crisis. This implies, for example, that a country j in period t with a reserve/import ratio of less than .67 is treated as being in crisis.[6] The text reports in detail the results for one such critical value, and in the final section explores the implications of changing that critical value.

Evidence of crisis and IMF participation in the data.

I consider the record of 90 developing countries on a quarterly basis during the period 1974:1 to 1992:4. The data for participation in IMF programs is drawn from IMF Annual Reports. The data on foreign exchange reserve holdings, imports and inflation rates are drawn from the International Financial Statistics CD-ROM.

The 90 developing countries in the sample can be differentiated by the degree of participation with the IMF. Figure 1 illustrates differentiation by the percent of the 1974:1 - 1992:4 time period the countries spent participating in IMF programs. Ten of the countries did not undertake an IMF program during the sample period. The modal range for participation was the 41-50 percent period, with 14 countries in that category. Four countries spent more than 70 percent of the period in IMF programs.

The various measures of crisis were chosen a priori as indicators of external imbalance, but they show little correlation among themselves or with participation in IMF programs. In Table 3 I report the Pearson correlation coefficients among the measure of participation in IMF programs I_{jt} and the various indicators of crisis.[7] While all indicators are significantly correlated

---

[6] For reserves growth, critical value indicating crisis was a greater than 10.2 percent reduction per period. For inflation, crisis was an outcome with greater than 4.2 percent quarterly inflation.

[7] The second statistic in each group is the test that the reported value is significantly different from zero, while the third number is the number of observations used in calculation.

with participation in IMF programs, the correlation coefficients are quite low. The largest, for the reserve/import ratio indicator, is 0.24. The inflation indicator C_{Ijt} has only an 8 percent positive correlation, while the reserve growth indicator C_{Gjt} actually presents a negative correlation with participation in IMF programs. The three indicators are not strongly positive associated with one another as well. For the illustrations that follow I choose C_{Rjt} to illustrate the transition observed in crisis.

The crisis transition matrix calculated from C_{Rjt} is given in Table 2. As is evident from the pattern of observations, there is a striking inertia to crisis. If there is a crisis observed in a quarter, the probability that the next quarter will be one of crisis as well is 85 percent. For quarters of non-crisis, the probability of continuing non-crisis is even higher at 92 percent. A chi-square test of non-random pattern indicates statistical significance to this inertial dynamic.[8] This pattern of significance is observed for more restrictive definitions of crisis as well.[9] Figure 2 divides the 86 countries into groups by the percentage of the sample that was spent in crisis as measured by C_{Rjt}. Fourteen countries had no periods of observed crisis by this definition, while four countries were in crisis in over 90 percent of the observed quarters.[10] The distribution, while not uniform, nevertheless includes 23 countries with more than half the observed period

---

[8] In this calculation, and those that follow, statistical significance is defined at the 95 percent level of confidence.

[9] It is clear from the statistics reported in Table A1 as well as from the definition of crisis that as this definition becomes more restrictive there is a greater likelihood that countries not in crisis will remain not in crisis. The off-diagonal elements of the transition matrix are roughly equal in observations, indicating a proportion of "new" crises and "solved" crises in the data. As the definition of crisis becomes more restrictive, the proportion of these non-inertial crises rises relative to total crises.

[10] These four countries were South Africa, Lesotho, Sudan and Romania. Since reserves are defined excluding gold, extensive gold reserves provide a partial explanation for South Africa's inclusion. I address that below in sensitivity analysis of the results.

characterized by crisis.

The arguments put forward by both proponents and critics of the IMF suggest that countries should participate in IMF programs in response to an episode of crisis. The hypothesis that crisis leads to participation implies a positive correlation that should be evident in the data. This correlation is most evident when the data are aggregated by country. If crisis and participation are linked, then for each country the share of the period 1973-1992 spent in crisis should be highly correlated with the share spent participating in IMF programs. This correlation is .85, and is both positive and significantly different from zero (but not from unity). A plot of these crisis/participation pairs by country is provided in Figure 3, with countries identified by three-digit IFS code.

Those countries above the 45^{o} line spent relatively more time in crisis than in IMF programs. South Africa (199) is the most striking outlier, while Senegal (722), Chad (628), Mali (678), Dominican Republic (243) and Lesotho (666) also have large positive differentials between crisis and participation. At the other extreme are the countries spending relatively more time in IMF programs than suggested by the crisis measure. These include the Philippines (566), Kenya (664), Togo (742), Mauritania (682), Uruguay (298), Argentina (213) and Mexico (273).[11]

Examination of individual data suggests a less powerful correlation between participation and crisis. As the evidence of Table 2 illustrates, when concurrent participation and crisis observations are examined the correlation coefficient is 0.24. This is significantly different from both zero and unity, as the associated probability value indicates. It is reasonable to interpret the

[11] Note that both groups include CFA zone members -- clearly, there is great deviation even within that zone.

hypothesized correlation between two consecutive periods, with the country entering a crisis in one period, beginning its negotiations with the IMF, and entering a program one or two periods later. These correlation coefficients are 0.25 and 0.26 respectively.

III. What is the contribution of IMF program participation to the entry into and exit from crisis?

While the crisis indicator has been defined, it is clear by the relatively low correlations reported in the previous section that this indicator may not define perfectly the periods of crisis facing the countries under observation. Suppose that there is an underlying crisis function c_{jt} and that it takes the form illustrated in equations (3), and the indicator function C_{jt} is defined as in (4).

$$c_{jxt} = c_{jt-1} + z_{j\{t\}}\delta + x_{j\{t\}}\beta_x + I_{j\{t-1\}}\gamma_x + \mu_{jx} + \nu_{jt} \qquad \text{for } c_{jt-1} < 0 \qquad (3a)$$

$$c_{jNt} = c_{jt-1} + z_{j\{t\}}\delta + x_{j\{t\}}\beta_N + I_{j\{t-1\}}\gamma_N + \mu_{jN} + \nu_{jt} \qquad \text{for } c_{jt-1} > 0 \qquad (3b)$$

The crisis indicator c_{jt} is equal to whichever of the two (c_{jxt} or c_{jNt}) is observed. Then

$$C_{jt} = \begin{cases} 1 & \text{for } c_{jt} + \epsilon_{jt} < 0 \\ 0 & \text{otherwise} \end{cases} \qquad (4)$$

The specification of equations (3) expands upon (1) by providing for the possibility that the impact of explanatory variables on crisis differ by whether country i was in crisis in the previous period. I provide an example in the appendix of foreign reserves policy by a developing-country

central bank that illustrates this type of regime shift in the crisis function. The specification of (4) provides for the possibility that the observed crisis indicator C_{jt} is an imperfect estimator of c_{jt}, and that the errors could be correlated between (3) and (4).

With an assumption on the distribution of ν_{jt} and ϵ_{jt} estimation of the parameters of interest is possible. I use two complementary techniques to derive these. The first derives the conditional probability of exit from crisis in any period t through use of probit estimation adjusted for the possibility of misidentification of crisis. The second exploits the fact that the magnitude of the conditional probability of exit from or entry to crisis is directly related to the expected duration of a period of crisis or of non-crisis. Using the statistical techniques of duration analysis, the impact of IMF programs on the duration of crisis and of non-crisis can be estimated.[12]

The model is specified to estimate the β_N and β_x – i.e., the effects common across countries of the various explanatory variables in the matrices $x_{jx\{t\}}$, – and the common IMF participation effects γ_N and γ_x. There is also country-specific and unobserved heterogeneity in the terms μ_{jx} and μ_{jN}, and this must be controlled for in estimation. I use a number of strategies to correct for this unobserved heterogeneity and will highlight them in turn as estimation results are discussed.

I consider two regime-specific explanatory variables $x_{j\{t\}}$ and two regime-specific measures of IMF participation $I_{j\{t-1\}}$ in the following estimation. The explanatory variables are indicators of past policy stance for the country i question. $R_{j\{t-1\}}$ is a measure of the lagged reserve/import

---

[12] As Lancaster (1990, ch. 2) illustrates, the conditional probability that the economy switches from one state to another can be modeled as a hazard function. Estimation of the Cox (1972) proportional hazard model for the duration of crisis or of non-crisis will provide estimates of the impact of current and cumulative IMF program participation on the conditional probability after controlling for initial conditions.

ratio. It will take one of two values in estimation – either R_{jt-1} or R_{j0}, with 0 indicating the period before the beginning of the recent episode of crisis or non-crisis. $\rho_{j\{t-1\}}$ is the real effective exchange rate, used both as ρ_{jt-1} and ρ_{j0} in what follows.[13] The variable I_{jt-1} indicates participation in an IMF program in the preceding period. The variable CI_{jt-1} measures the quantity of periods from the previous twelve that country j participated in IMF programs. The β and γ vectors thus have 2 elements apiece in the reported results, with their values potentially varying by whether the period is one of crisis or non-crisis.

$$x_{j\{t\}}\beta_k = \beta_{rk}R_{jt-1} + \beta_{\rho k}\rho_{jt-1} \quad \text{or} \quad x_{j\{t\}}\beta_k = \beta_{rk}R_{j0} + \beta_{\rho k}\rho_{j0} \tag{5a}$$

$$I_{j\{t-1\}}\gamma_k = \gamma_{1k}I_{jt-1} + \gamma_{2k}CI_{jt-1} \quad \text{for } k=N,x \tag{5b}$$

There are as well factors common to both regimes captured by $z_{j\{t\}}\delta$ in (3). These include a "normal" probability of transition and country and time specific effects.

The hypothesis tests of this paper can be nested in this formulation. The null hypothesis of no impact of participation in IMF programs on crisis will be that the elements of γ_k (k=N,x) are insignificantly different from zero. [14] The first alternative hypothesis is consistent with the dynamic found in the IMF Articles of Agreement: γ_x will be positive and γ_N will be negative with participation. Thus the probability of escaping crisis is increased while the probability of shifting from non-crisis to crisis is reduced. The second alternative hypothesis is the critics' position: that

[13] The real effective exchange rate is available from the IMF. It is normalized in this analysis by country by subtracting the mean of the series for the longest period available and dividing by the standard deviation of that series for that same period.

[14] This hypothesis is in fact a venerable critique of IMF activity as well, as found in the charge Spraos (1986) levels in his title: IMF conditionality is "ineffectual".

γ_x will be negative while γ_N will be positive. Under this hypothesis the transition from crisis to non-crisis is less likely while the transition from non-crisis to crisis is more likely due to the inappropriateness of the policy advice of the IMF.[15]

Statistical techniques.

There are four possible transitions from period t-1 to period t, with conditional probabilities of entry into and exit from crisis defined as:

$$\begin{aligned} p_{jNt} &= P(c_{jt} = 1 \mid c_{jt-1} = 0) \\ p_{jxt} &= P(c_{jt} = 0 \mid c_{jt-1} = 1) \\ (1- p_{jNt}) &= P(c_{jt} = 0 \mid c_{jt-1} = 0) \\ (1- p_{jxt}) &= P(c_{jt} = 1 \mid c_{jt-1} = 1) \end{aligned} \tag{6}$$

If C_{jt-1} were a perfect indicator of c_{jt-1}, then equations (3) could be estimated through partition of the dataset by value of C_{jt-1}. However, the selection bias involved in the transition from one period to the next implies that the ν_{jt} in equations (3) are not zero-mean in nature. The imperfect observation of crisis and the possibility that ϵ_{jt-1} will be correlated with ν_{jt} requires that this estimation include the correction M_{jkt-1} defined in equations (7). If the covariance between ν_{jt} and ϵ_{jt-1} is defined σ_{xc} for C_{jt-1}=1 and σ_{Nc} otherwise, and the probability and cumulative distribution functions of ν_{jt} are $\phi(.)$ and $\Phi(.)$, respectively, then the expected values can be defined:

$$\begin{aligned} E(\nu_{jt} \mid c_{jt-1} = 1) &= - \sigma_{xc} M_{jxt-1} \\ E(\nu_{jt} \mid c_{jt-1} = 0) &= \sigma_{Nc} M_{jNt-1} \end{aligned} \tag{7}$$

---

[15] There are also transition probabilities for participation in IMF programs. In a companion paper, I examine the impact of country-specific performance or shocks on the decision to participate; these may be contemporaneous, or may be dependent upon observed behavior in periods before t. The existence of crisis, or crisis over a number of preceding periods, is then jointly determined with participation in IMF programs in a dynamic system.

$$M_{jxt-1} = \phi(\nu_{jt-1})/\Phi(\nu_{jt-1})$$

$$M_{jNt-1} = \phi(\nu_{jt-1})/[1-\Phi(\nu_{jt-1})]$$

The equations determining c_{jxt} and c_{jNt} can be rewritten as:

$$c_{jxt} = z_{t\{t\}}\delta + x_{j\{t\}}\beta_x + I_{j\{t-1\}}\gamma_x + \mu_{jx} - \sigma_{xc}\, M_{jxt-1} + \epsilon_{jt} \qquad \text{for } C_{jt-1} = 1 \qquad (8a)$$

$$c_{jNt} = z_{t\{t\}}\delta + x_{j\{t\}}\beta_N + I_{j\{t-1\}}\gamma_N + \mu_{jN} + \sigma_{Nc}\, M_{jNt-1} + \epsilon_{jt} \qquad \text{for } C_{jt-1} = 0 \qquad (8b)$$

ϵ_{jt} will have zero conditional mean in each equation.

M_{jkt-1} is correlated with the error in equations (8), but a two-stage estimator can be derived. In the first stage a predictor m_{jkt-1} can be derived for M_{jkt-1} from probit estimation over the explanatory variables that are invariant to regime.

$$c^{*}_{jt-1} = z_{j\{t-1\}}\delta + \nu^{*}_{jt-1} \tag{9}$$

$$C_{jt-1} = \begin{cases} 1 & \text{for } c^{*}_{jt-1} > 0 \\ 0 & \text{otherwise} \end{cases} \tag{10}$$

These explanatory variables will be the observation on crisis in the preceding period as well as country-specific and time-specific dummy variables. The m_{jkt-1} are then derived from the estimates of ν^{*}_{jt-1}. They are used as explanatory variables in estimation of equations (8) over the two subsamples defined by the value of C_{jt-1}. The standard errors of the estimates must then be adjusted for the use of this instrumental variable.

Probit analysis.

The conditional probabilities p_{jxt} and p_{jNt} of equations (6) can be derived under assumption of normality of errors through two-stage probit analysis of the indicator variable C_{jt}. Panel estimation occurs over the 6840 pooled observations, although the absence of observations on some variables (e.g., the real exchange rate) and truncation of the sample using only the observations in the period 1976-1992 to permit use of CI_{jt-1} leads in most estimation results to smaller sample sizes. In the first stage, as indicated above, the equation system (9) and (10) is estimated with country-specific and time-specific dummy variables and with the lagged observed indicator of crisis.[16] This estimation generates the m_{jkt-1} estimates used in the following second-stage probit estimates of equations (8) and (4).

The four columns of Table 4 report the results of probit analyses of the determinants of

[16] This first stage of estimation is not reported in the paper, but the results are available from the author on request.

the conditional probabilities of exit from crisis p_{jxt} and entry into crisis p_{Nt} for the crisis indicator C_{Rjt}, while similar tables are reported in the appendix for the other two crisis indicators. The first and third columns of figures (denoted "Simple") present results when unobserved country-specific heterogeneity is not modeled, while the second and fourth columns of figures provide results contingent upon fixed-effect correction for country-specific heterogeneity. The first two columns use as explanatory variables the participation in an IMF program in a previous period (I_{jt-1}), the cumulative participation in IMF programs over the previous 12 periods (IC_{jt-1}), the real exchange rate observed in the year prior to the beginning of crisis (ρ_{j0}) and the reserve-import ratio observed in the year prior to the beginning of crisis (R_{j0}).[17] Other things equal, I anticipate that increases in $R_{j\{t\}}$ and in $\rho_{jt\{t\}}$ (a depreciation) will lead to increased conditional probabilities of exit from crisis. Single asterisks after the coefficient estimates indicate that the estimate is significantly different from zero at the 90 percent level of confidence, while a double asterisk indicates significance at the 95 percent level of confidence.

Comparison of the first and second columns of figures illuminates a general conclusion from this research about the appropriate statistical methodology. The estimation strategy in the first column does not control for unobserved heterogeneity in the occurrence of crisis, while the strategy of the second column does so through inclusion of fixed effects. There is significant evidence in the first column of selectivity bias (as indicated by the coefficient of m_{jxt-1}) as well as a significant effect of depreciation of the real exchange rate and a higher R_{j0} on the probability of exit. The message for IMF program participation is not good – the direct effect through

[17] Similar analyses were completed using a one-period lagged value of R_{jt-1} and ρ_{jt-1}. These give similar results for the IMF-related variables, and are available on request.

participation is positive but insignificantly different from zero at 0.21, while the cumulative effect as measured by the coefficient of IC_{jt-1} is negative and significantly different from zero at the 90 percent level of confidence. Correction of country-specific heterogeneity, however, yields different conclusions. The evidence of significant selection bias disappears. The effect of real depreciation in exiting crisis remains unchanged. The direct impact of IMF participation on the probability of exit from crisis is positive and significant at the 95 percent level of confidence, while the evidence of a negative cumulative effect remains but is insignificantly different from zero.

The evidence of the probability of entering crisis has the same methodological lesson, and so I focus upon the column reporting the results with correction for unobserved heterogeneity. Unsurprisingly, it is less likely that a country will enter crisis if it begins with larger reserve holdings. An initial real depreciation also reduces the probability of entering crisis, although insignificantly. Participation in IMF programs reduces the probability of entry into crisis directly, as indicated by the -0.40 coefficient, while the cumulative effect of long-term participation in IMF programs, represented by the 0.65 coefficient, is to increase the conditional probability of entry into crisis. The relative importance of these effects will be illustrated in a simulation in a later section; however, since IC_{jt-1} takes values between 0 and 1 it is evident that the positive direct effect of participation in IMF programs during non-crisis years will be more-than-completely counterbalanced by the cumulative effect of participation for values of IC_{jt-1} greater than (.40/.65).

Survival analysis.

The probability of exit from crisis can also be inferred from examination of the duration

observed for crisis episodes in the data. The probability of exit in period t conditional on crisis existing at the beginning of the period can be expressed as the ratio of the unconditional probability of exit in that period to the probability that the crisis continued to the beginning of period t.[18] Thus, knowledge of the duration of crisis provides insight into the conditional probability of exit from crisis. Duration of the absence of crisis similarly provides insight into the conditional probability of entry into crisis.

Equations (6) can be represented by standard formulations of hazard functions in duration analysis. However, there are once again two potential difficulties in estimation. First, there is the possibility of unobserved heterogeneity in the conditional probabilities associated with the country in crisis. Second, there is the possibility that the beginning of crisis (or non-crisis) is not precisely measured. Therefore, I introduce corrections for country-specific heterogeneity and for selection bias in Tables 5 and 6.

Table 5 reports the coefficient estimates and associated statistics for the proportional-hazards model in considering crisis indicator C_{Rit}. In the first column the μ_{jx} is treated as a random effect with identical mean but heteroskedastic variance across countries. Coefficient estimates for the four explanatory variables are reported, each followed in the column by the standard error of the estimate and the χ^2 statistic of the Wald test that the coefficient is equal to zero. A White correction for unspecified heteroskedasticity is used for the standard errors. The final statistic for each coefficient is the risk ratio, indicating the ratio of the value of the conditional probability when the explanatory variable is evaluated at unity and at zero. For the coefficient of the dummy

[18] If the probability distribution function for country j to exit during period t is f_{jt} and the probability of crisis continuing up until t is S_{jt} – also called the survivor function – then $p_{jt} = f_{jt}/S_{jt}$. Lancaster (1990) has a useful development of these probability concepts.

variable I_{jt-1}, for example, the risk ratio of 1.76 indicates that a country in an IMF program has a 76 percent higher conditional probability of exiting in each period of crisis than a country without a program. For IC_{jt-1}, the risk ratio of 0.91 indicates that for each additional period of the preceding twelve spent participating in an IMF program, the country has a nine percent lower probability of exit.

The other two coefficients in this estimation have sensible interpretations and are significantly different from zero. The larger the R_{j0} when crisis begins the larger the conditional probability of exit, with an increase in this ratio by 0.1 leading to an increase in the conditional probability by 2.3 percent. The larger (i.e., more depreciated) the ρ_{j0} observed in the year prior to the beginning of crisis the higher the conditional probability of exit in each period, other things equal: a 0.1 unit increase in the normalized real exchange rate leads to an increase in the conditional probability by 1.4 percent. The selection bias correction also proves to be significant.

While these estimation results take a form consistent with theory, it will be useful to address explicitly the unobserved heterogeneity. In the second column of results this is done through stratification estimation (i.e., inclusion of separate μ_{jx} terms for each of the countries included in the sample). Coefficient estimates on the explanatory regressors are assumed once again to be equal across countries. This fixed-effect estimation improves the overall explanatory power of the estimation, but at the cost of greatly increasing the degrees of freedom used in the statistical exercise.[19] The coefficient estimates indicate a positive and significant effect (.78) of

[19] There are 143 observations in this data set from 62 countries used in the estimation. In the first column these 143 observations were used to identify four estimated coefficients. In the second column there were in effect 66 coefficients estimated, with a resulting loss of efficiency (noted in the standard errors of estimates). For example, in 15 of the countries there was only one exit from crisis. Only in the remaining 47 countries are there sufficient exits from crisis to provide degrees of freedom in estimation.

participation in IMF programs on the probability of exit, while the cumulative effect of IMF programs disappears. Other statistical characteristics of the results suggest caution in interpreting these, however, especially the negative and significant effect of reserve holdings on the probability of exit.

The stratification results of column two control for the heterogeneity across countries through the introduction of fixed-effect terms, and apparently addresses selection-bias concerns as well. Alternatively, random-effect terms can be introduced that model explicitly the heteroskedasticity across countries. This approach requires assumption of a distribution of the random-effect terms; here I follow Hougaard (1986) and Shu and Klein (forthcoming) in assuming that the random effects are independent and identically distributed positive stable variates.[20] The single parameter of this distribution, θ, lies between 0 and 1 and indexes the degree of heterogeneity across groups relative to within the group. With $\theta=1$ the sample is drawn from an independent distribution with no within-group dependence. Imposition of $\theta=1$ yields the estimation results of the first column. For $\theta< 1$ there is less variation for observations from a specific country than across countries.

The final column of Table 5 indicates the results of estimation in this random-effects framework. Examination of the coefficient estimates of θ at the bottom of the table indicates that both hypotheses of independence and of purely fixed effects can be rejected; the χ^2 statistic is reported for the test of difference from unity, but is also evidently significantly different from zero in each case. The direct impact of participation is similar in size to that of the stratified sample

[20] As Lancaster (1990) points out, this distribution has the property that the joint distribution of ν_{jxt} and ϵ_{jx} will preserve the proportional-hazards property of the fixed-effect estimation.

and is significantly different from zero; the cumulative effect of participation is negative but insignificant. The explanatory variables R_{j0} and ρ_{j0} have insignificant sign. The correction for selection bias is once again significant.

The preponderance of evidence in these estimation results indicates a significant and positive impact of participation in an IMF program on the conditional probability of exit from crisis -- this effect is quite large in quantitative terms as well. The cumulative impact of prior participation is negative, but is significant in only one of the three estimation techniques.

Survival analysis allows for the possibility of conditional probability of exit dependent upon the time spent in crisis (i.e., non-constant $\phi(s)$), and in this instance the derived probability is quite non-linear in the duration of the crisis. Figure 4 illustrates the estimate of the baseline probability of exit derived from the proportional-hazards estimation stratified by country.[21] The baseline probability begins with a large value and then falls. At about the 15^{th} quarter the probability of exit begins to rise through the 30^{th} quarter, and then declines again.

Table 6 indicates a similar derivation for the probability of entry into crisis (or alternatively, an exit from non-crisis). Both selection bias and unobserved heterogeneity are significant features of the data. The stratified sample and frailty measure results give roughly similar results, with insignificant direct and cumulative effects of participation in IMF programs on this conditional probability. The sign of the reserve holdings is the expected sign in these results, but is insignificantly different from zero. The sign and significance of the real exchange rate is

[21] The proportional-hazards estimation does not explicitly model the baseline probability. The function reported in Figure 4 is derived from the implied average conditional probability for each duration through a kernel smoothing process with bandwidth of 8.4 periods. The initial and final quarters are then excluded through the smoothing process.

surprising -- it indicates that a real depreciation, other things equal, will increase the probability of entry into crisis. This may be a consequence of the increased volatility attached to the external orientation implied by the depreciated real exchange rate; such volatility will make crises more likely. However, further study is necessary before this conclusion can be decisively drawn.

Simulation results.

The dynamic consequences of participation in IMF programs are best examined in the context of a simulation exercise. In Figures 5 and 6 I illustrate the outcome of a 42-period simulation drawn from a longer simulation to illustrate the various effects of IMF programs. The variables for R_{j0}, ρ_{j0} and ν_{jt} were generated in random fashion, with the country-specific effect assumed of intermediate size (so that the country will be at times in crisis and at times not). The country is assumed to enter an IMF program in the period in which it enters into crisis, and the cumulative total of IMF participation is calculated accordingly. At the beginning of the simulation the country was out of crisis, and had been so for at least 12 periods.

The country's performance is illustrated in Figure 5 by the line labelled "with IMF programs". The random shocks buffeting the economy were negative enough to lead to seven distinct episodes of crisis. In each of these crisis periods the participation in IMF programs contributed to a tendency toward non-crisis, although as IMF participation continued that tendency was attenuated. Figure 6 illustrates the numerous switches from IMF participation to non-participation, as well as the index of cumulative participation.

The country's performance without IMF participation is illustrated by the lower graphic in Figure 5. For this simulation the random shocks are assumed the same, as is the switching of behavioral functions based upon crisis or non-crisis in the previous period. However, the country

is not given the opportunity to participate in IMF programs. The country moves into crisis only twice, and descends deeply into crisis on the second pass.

A third "comparative dynamic" tracks the behavior of this economy if, while facing similar shocks, it was not saddled with the negative consequences of cumulative IMF participation. The upper graphic in Figure 5 illustrates this outcome. There are three episodes of crisis early in the simulation period, and the direct impact of IMF participation is sufficient to drive the economy into a sustained period of non-crisis. There is less switching, clearly, as the negative effect of cumulative IMF participation does not hold back the economy's performance.

IV. Tests for robustness.

Use of alternative explanatory variables.

The explanatory variables used in the definition of conditional probability were in part dictated by the frequency and cross-country available of the data. However, the use of R_{j0} and ρ_{j0} was an effort to introduce comparability into the probit and survival analyses. When the probit analysis is redone using R_{jt-1} and ρ_{jt-1}, the estimation results (not reported here) yield nearly identical conclusions on the hypothesis tests concerning participation in IMF programs. The reserve-import ratio becomes a quantitatively more important explanatory variable, while the quantitative effect of the real exchange rate on the conditional probability of exit is reduced.

Use of alternative measures of crisis.

The conclusions of this paper are clearly predicated upon the choice of indicator of economic crisis. As noted above, I investigated two alternative measures in the course of this research: C_{jGt} and C_{jIt}. Tables A4.1 and A4.2 report the results of examination of conditional

probabilities from probit estimation for these two measures, while Tables A5.1 and A5.2 provide results analogous to Table 5.

As noted in Table 3, these other measures of crisis are less positively correlated with IMF participation. Therefore, I anticipate less striking results when the analysis is redone using these. The tests when the indicator R_{Git} is used for crisis are reported in Table A4.1. The direct impact of IMF programs on the probability of exit from and entry to crisis takes the correct sign in all cases, and is significantly different from zero for the fixed-effects estimation. The cumulative impact of IMF programs proves to have the same, but insignificant, effect in this formulation. The lesson from this crisis indicator is a simple one – the probability of exiting a reserve-growth crisis declines with an increase in the initial stock of reserves. If countries can sustain negative reserve growth longer then they will do so.

In the analysis of the inflation indicator of crisis C_{Iit} in Table A4.2 there is once again the correct signs to the direct impact of participation in IMF programs on the conditional probabilities, although in no case is the effect significant. The cumulative effects take signs opposite to those of the direct effect, as noted also in Table 4, but these are also insignificant.

The probability of exit from crisis can be examined for these indicators of crisis using survival analysis, and these results are reported in Tables A5.1 and A5.2. When crisis is defined as in R_{jGt}, the direct effect of participation is always positive and significant, while the cumulative effects are negative but insignificant. When survival analysis is used with crisis indicator R_{jIt}, the contribution of IMF programs to the conditional probability is small and insignificant for both direct and cumulative measures.

The ratio of reserves to imports used in this paper can be criticized on two grounds. First,

the stock of gold is excluded from the measure of reserves. Second, the appropriate measure of uses of reserves may not be the value of merchandise imports but the value of all debits in the balance of trade in goods and services, including debt service and other service liabilities as well as merchandise imports. In response to these shortcomings I redid the preceding analysis with two additional reserve ratios. The initial reserve/imports ratio is denoted h_{it}. The first new ratio, denoted hg_{it}, has reserves inclusive of gold as numerator and merchandise imports as denominator. The second new ratio, denoted hh_{it}, is the ratio of reserves inclusive of gold to total current-account debits other than transfers. The data are drawn from the IMF International Financial Statistics.

These new series are available for only a subset of the sample used above. Missing values for gold holdings eliminated 22 countries and 788 observations from the sample, while missing observations for the current-account debits eliminated another 2545 observations. The samples for the three analyses are thus not in general the same. The correlations (ρ) of h_{jt} with hg_{jt} and hh_{jt} are strong but not perfect, with $\rho(h_{jt},hg_{jt}) = 0.93$, $\rho(h_{jt},hh_{jt}) = 0.80$, and $\rho(hg_{jt},hh_{jt}) = 0.79$. While $\rho(h_{jt},I_{jt}) = -0.22$, $= -0.16$ and $\rho(hh_{jt},I_{jt}) = -0.15$ over the truncated samples. While some countries have been excluded by data limitations, the pattern of crisis relative to IMF participation of the larger sample is replicated. South Africa, to take the most extreme case, remains much more often in crisis than in an IMF program over the period. The results of probit estimation replicate those of Table 4.

The definition of IMF programs used in the preceding analysis is an inclusive one. First tranche drawings are combined with stand-by, EFF, SAF and ESAF programs. While all such programs are logical responses to crisis, they differ in important ways in the conditions imposed

upon the participating country. I test the sensitivity of the results above by creating an alternative measure of IMF programs that excludes first-tranche drawings.[22] The determinants of entry into and exit from crisis are also much as in the earlier discussion.

V. Conclusions and extensions.

While a casual examination of the data suggests that participating in IMF programs can be hazardous to a country's economic health, more detailed examination suggests a more nuanced view. First, there is significant evidence in two different indicators of economic crisis that participation in IMF programs leads to an increased probability that the country will leave crisis (or alternatively, that the crisis will be of shorter duration). Second, the impact of continuing reliance upon participation in IMF programs lessens this effect, even when other country-specific characteristics have been controlled for.

When the country is not in crisis but nevertheless participating in IMF programs, then the story is less clear. The effects of participation on the probability of entering crisis are insignificant, but there is some evidence that participation in IMF programs in a given year of non-crisis makes more likely the return to a crisis state in the following year. (This may be an effect of "pre-emptive" participation. Persistent IMF program participation in the previous three years, though has a weak effect in extending the period of non-crisis.

A puzzle remains, however. All measures of crisis are much less than perfectly correlated with participation in IMF programs. This makes it important to examine the other half of this

---

[22] I also redefine the cut-off value h^{crit} to include only 30.6 percent of observations, since in this analysis the "first-tranche" periods will be treated as non-program periods.

question – i.e., what are the dynamic determinants of IMF participation? This is an area in which my Political Science colleague Tim McKeown and I are presently working.

These conclusions are of course dependent upon (1) the use of the crisis measures chosen, the choice of critical values, and (3) the definition of IMF programs. I have used a number of specifications to check for robustness, but there is certainly more room for such activity.

Bibliography

Agenor, P-R. and P. Montiel: Development Macroeconomics. Princeton, NJ: Princeton University Press, 1996.

Bird, G.: "The Myths and Realities of IMF Lending", World Economy 15, 1994, pp. 759-778.

Black, S.: "International Money and International Monetary Arrangements", in R. Jones and P. Kenen, eds.: Handbook of International Economics, volume 2. Amsterdam, NL: North-Holland, 1985.

Conway, P.: "IMF Lending Programs: Participation and Impact", Journal of Development Economics 45, 1994, pp. 365-391.

Conway, P.: "Recidivism in IMF Programs and the Economics of Crisis: Evidence from the period 1973-1992", processed, 1997.

Conway, P. and T. Mroz: "IMF Programs and Foreign-Exchange Reserve Holding: Evidence on the Joint Distribution", processed, 1998.

DeGroot, M.: Probability and Statistics. Reading, MA: Addison Wesley, 1973.

Diette, T.: "Does Prior Participation Affect the Duration of IMF Spells?" UNC Masters Thesis, 1999.

Economist, The: "New Illness, Same Old Medicine", 13 December 1997, pp. 65-66.

Feldstein, M.: "Focusing on Crisis Management", op-ed piece, Wall Street Journal, 6 October 1998.

Fischer, S.: "Applied Economics in Action: IMF Programs", American Economic Review (Paper and Proceedings) 87, 1997, pp. 23-27.

Hamilton, J.: Time Series Analysis. Princeton, NJ: Princeton University Press, 1994.

Heckman, J. and B. Singer: "A Method for Minimizing the Impact of Distributional Assumptions in Econometric Models for Duration Data", Econometrica 52/2, 1984.

Hougaard, P.: "Survival Models for Heterogeneous Populations Derived from Stable Distributions", Biometrika 73/2, 1986, pp. 387-396.

Hougaard, P.: "A Class of Multivariate Failure Time Models", Biometrika 73/3, 1986, pp. 671-678.

Kenen, P.: Financing, Adjustment and the International Monetary Fund. Washington, DC: Brookings Books, 1986.

Khan, M.: "The Macroeconomic Effects of Fund-Supported Stabilization Programs", IMF Staff Papers 37/2, 1990, pp. 195-231.

Killick, T.: The Quest for Economic Stabilization: The IMF and the Third World. London, UK: Heinemann Educational Books, 1984.

Killick, T.: IMF Programmes in Developing Countries: Design and Impact. London, UK: Routledge, 1995.

Killick, T., M. Malik and M. Manuel: "What Can We Know About the Effects of IMF Programmes?", World Economy 15, 1995, pp. 575-597.

Knight, M. and J. Santaella: "Economic Determinants of IMF Financial Arrangements", Journal of Development Economics, 1998, forthcoming.

Lancaster, T.: The Econometric Analysis of Transition Data. Cambridge, UK: Cambridge University Press, 1990.

Maddala, G.: Limited Dependent and Qualitative Variables in Econometrics. Cambridge, UK: Cambridge University Press, 1983.

Oakes, D.: "Bivariate Survival Models induced by Frailties", Journal of the American Statistical Association, 84/406, 1989, pp. 487-493.

Obstfeld, M. and K. Rogoff: Foundations of International Macroeconomics. Cambridge, MA: MIT Press, 1996.

Rosett, C.: "The World's Poor Pay the Price for the IMF's Failures", op-ed piece, Wall Street Journal, 22 April 1999.

Sachs, J.: "The Wrong Medicine for Asia", op-ed piece, New York Times, 3 November 1997.

Sachs, J.: "Rule of the Ruble", op-ed piece, New York Times, 4 June 1998.

Shu, Y. and J. Klein: "A SAS Macro for the Positive Stable Frailty Model", American Statistical Association, Proceeding of the Statistical Computing Section, 1999.

Spraos, J.: IMF Conditionality: Ineffectual, Misguided, Mistargeted. Princeton Essays in International Finance 166. Princeton, NJ: International Finance Section, 1986.

Figure 1
Participation in IMF programs: by country

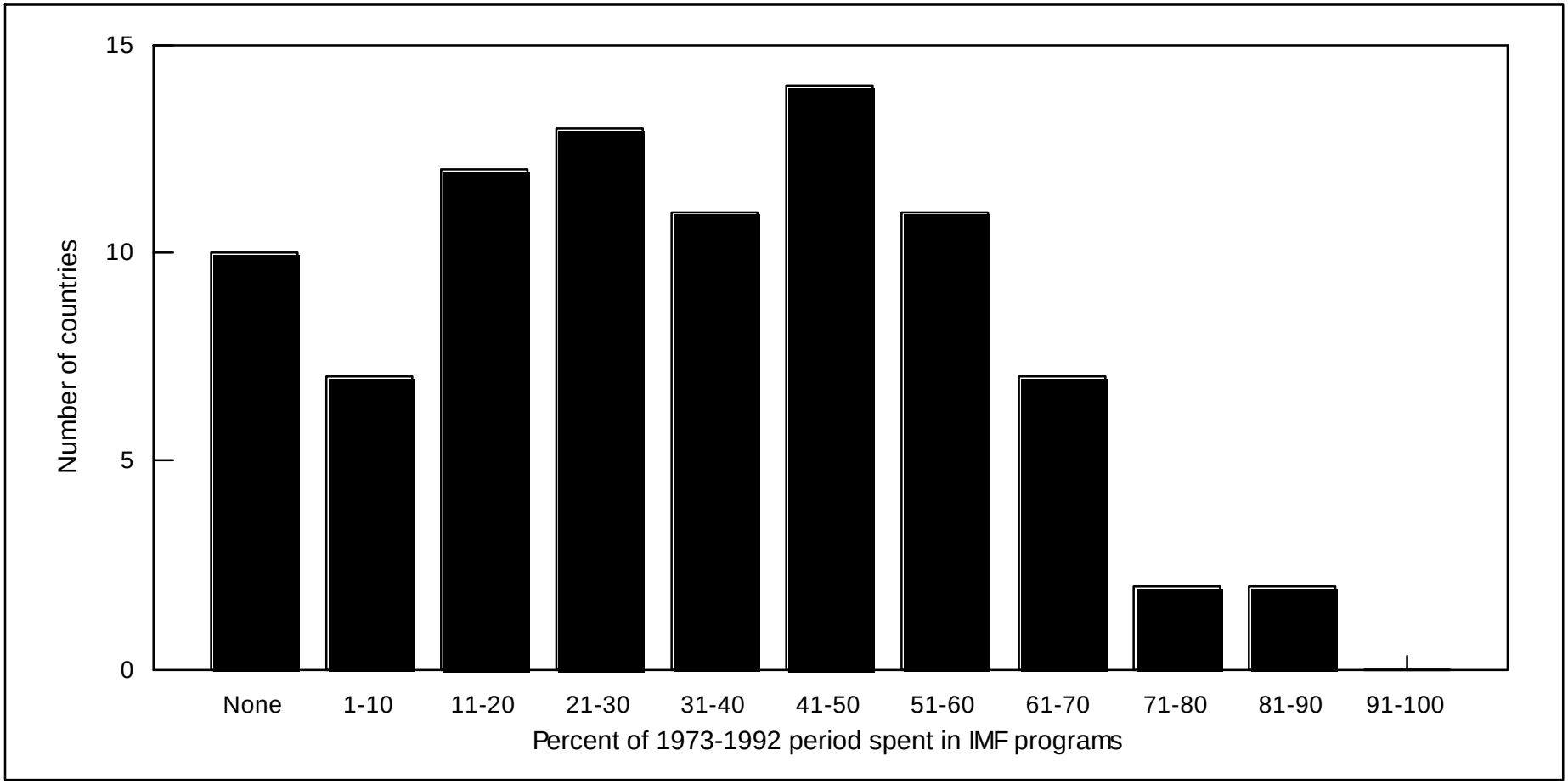

Figure 2
Percent of Sample Period spent in Economic Crisis: by country

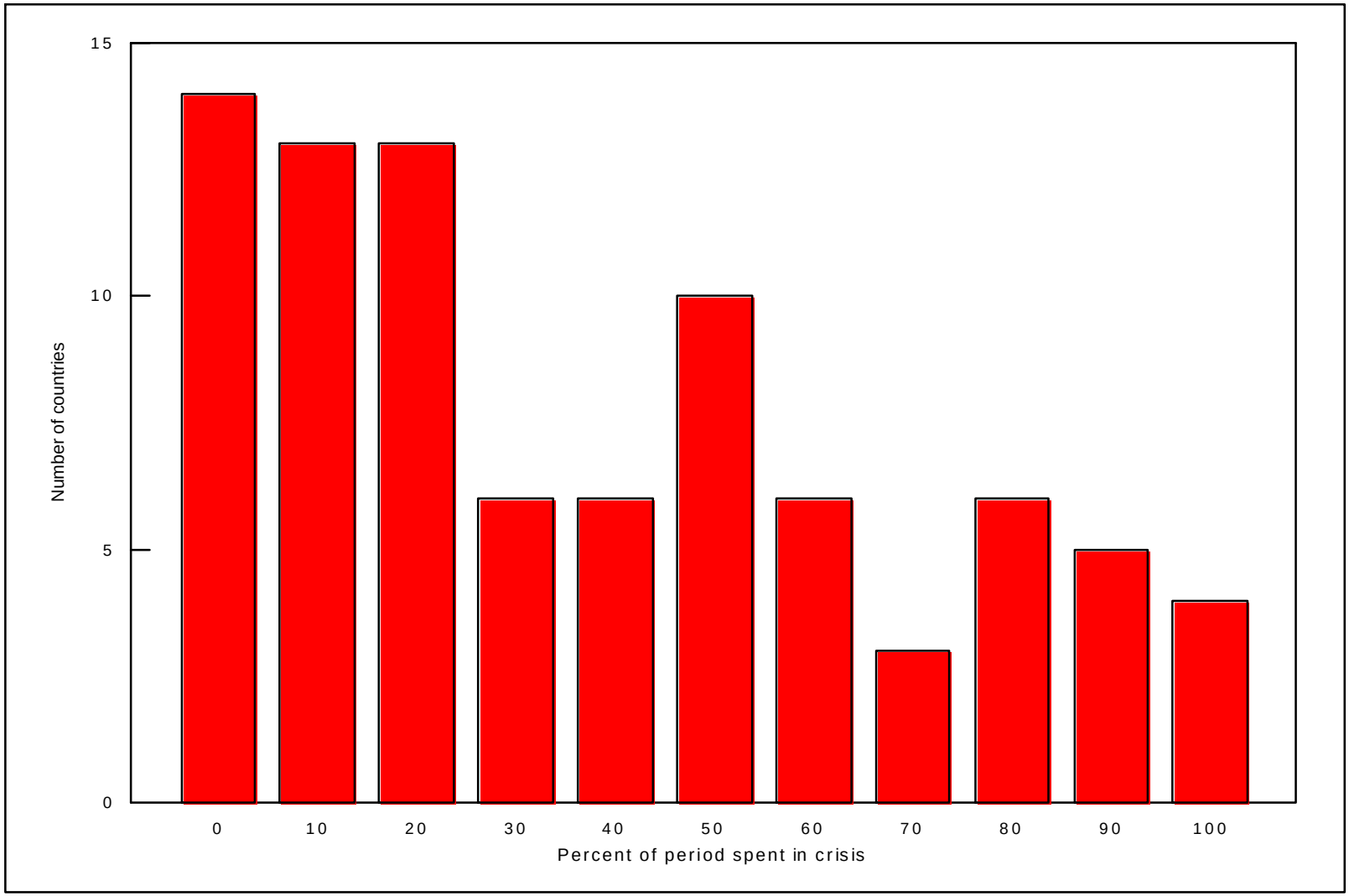

Figure 3
Crisis and participation in IMF programs: by country

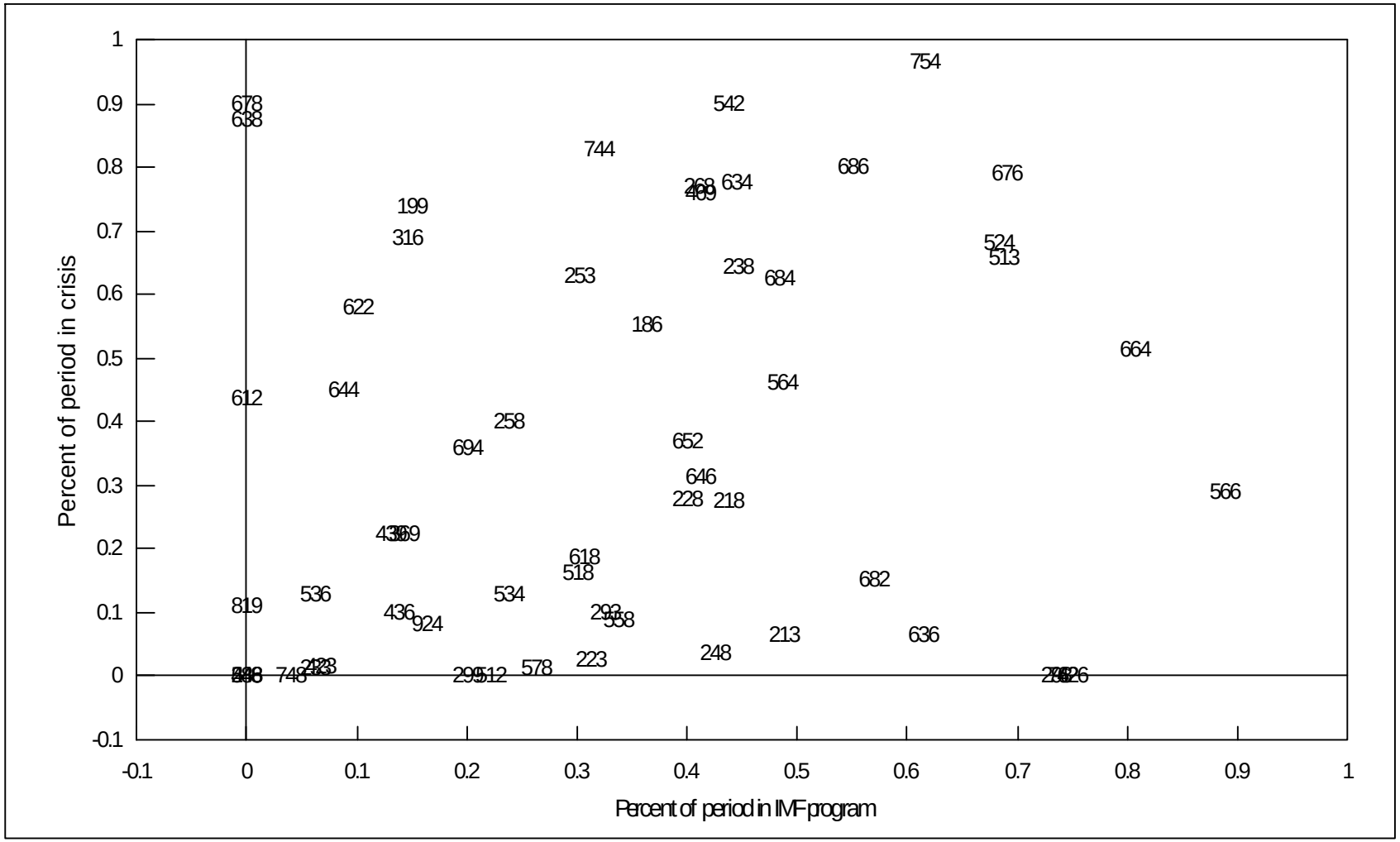

Figure 4

Baseline Conditional Probability of Exit From Crisis

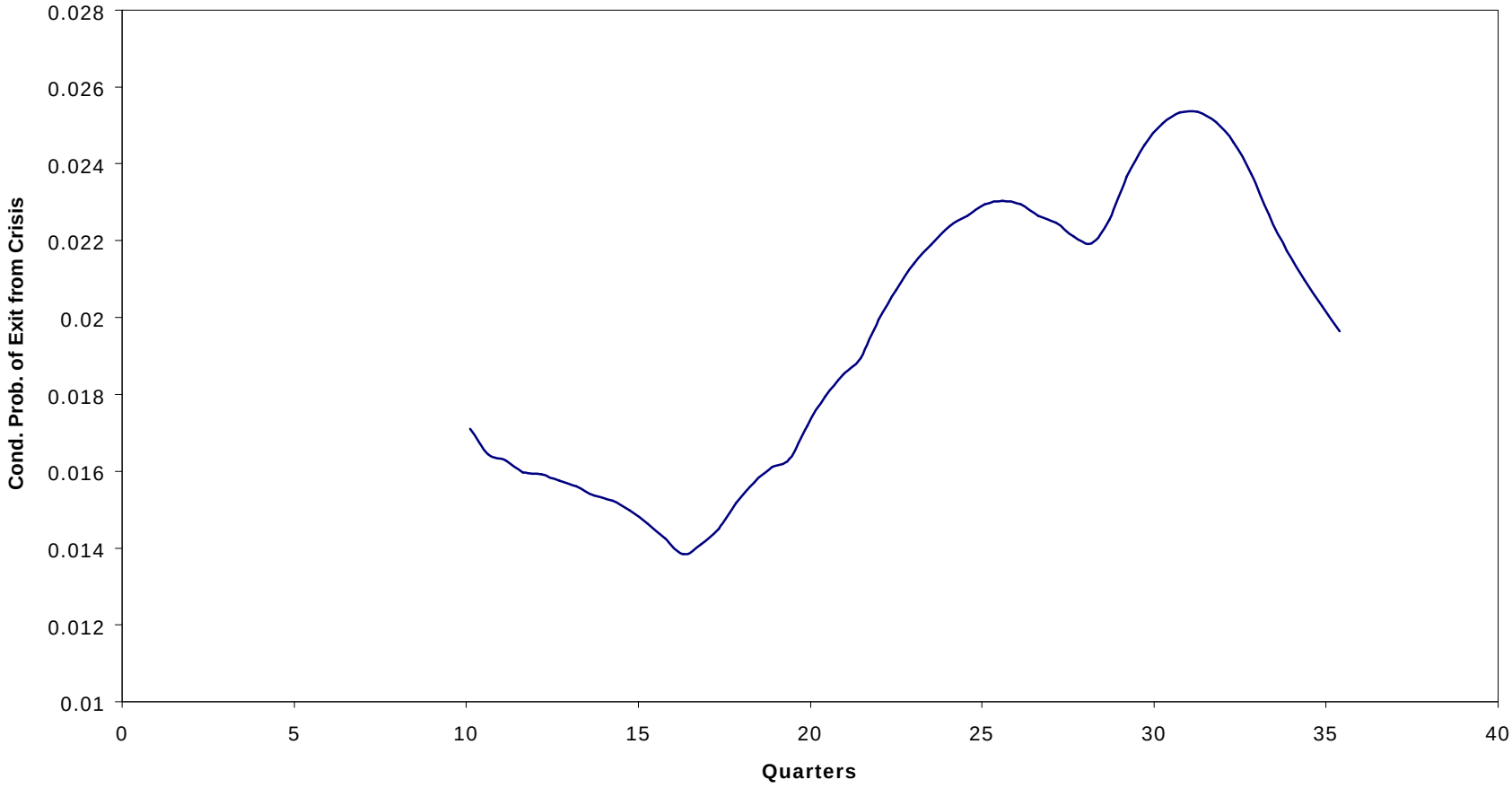

Figure 5
Implied Crisis Generated by Switching Regression

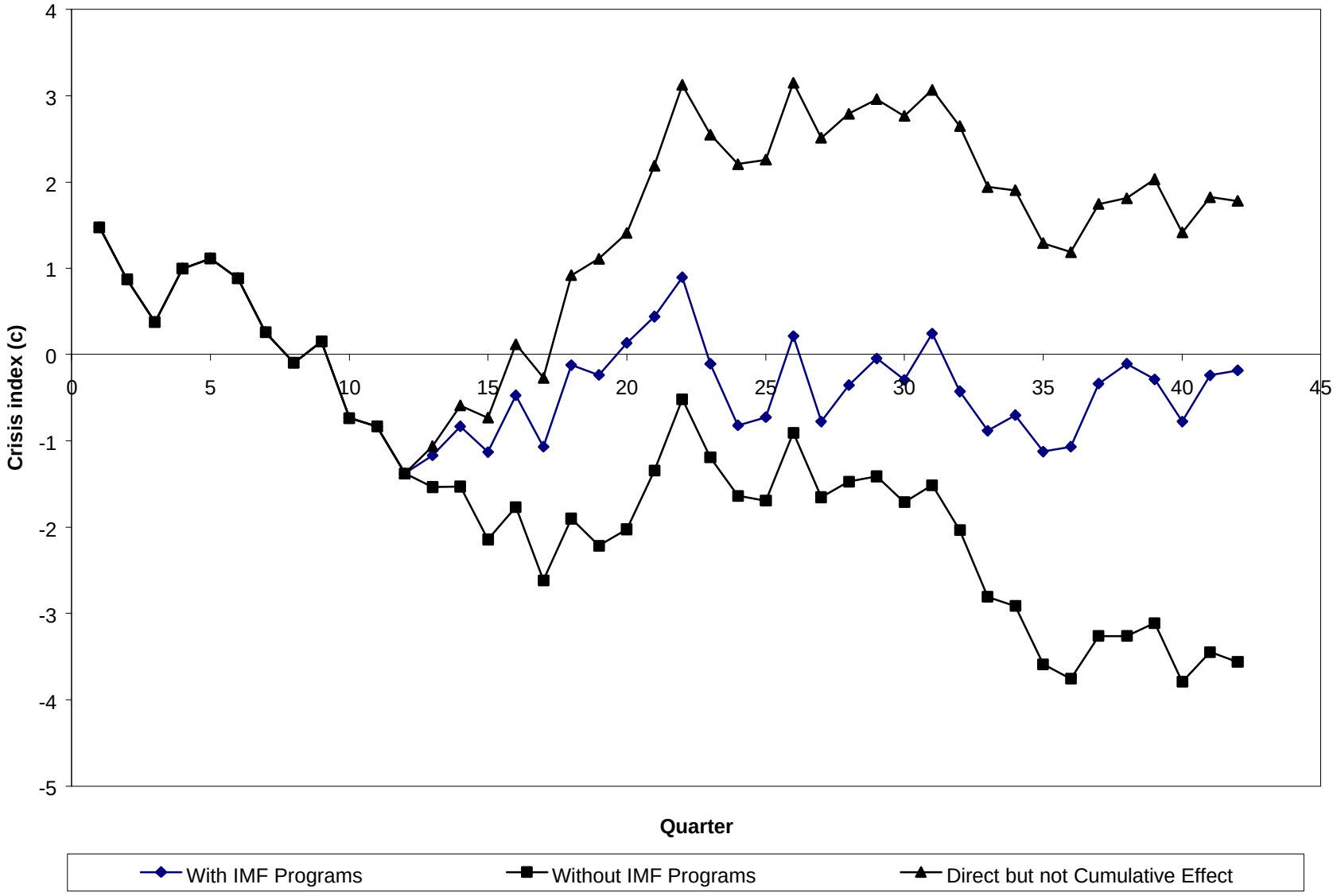

Figure 6
Simulation Results: Direct and Cumulative IMF Participation

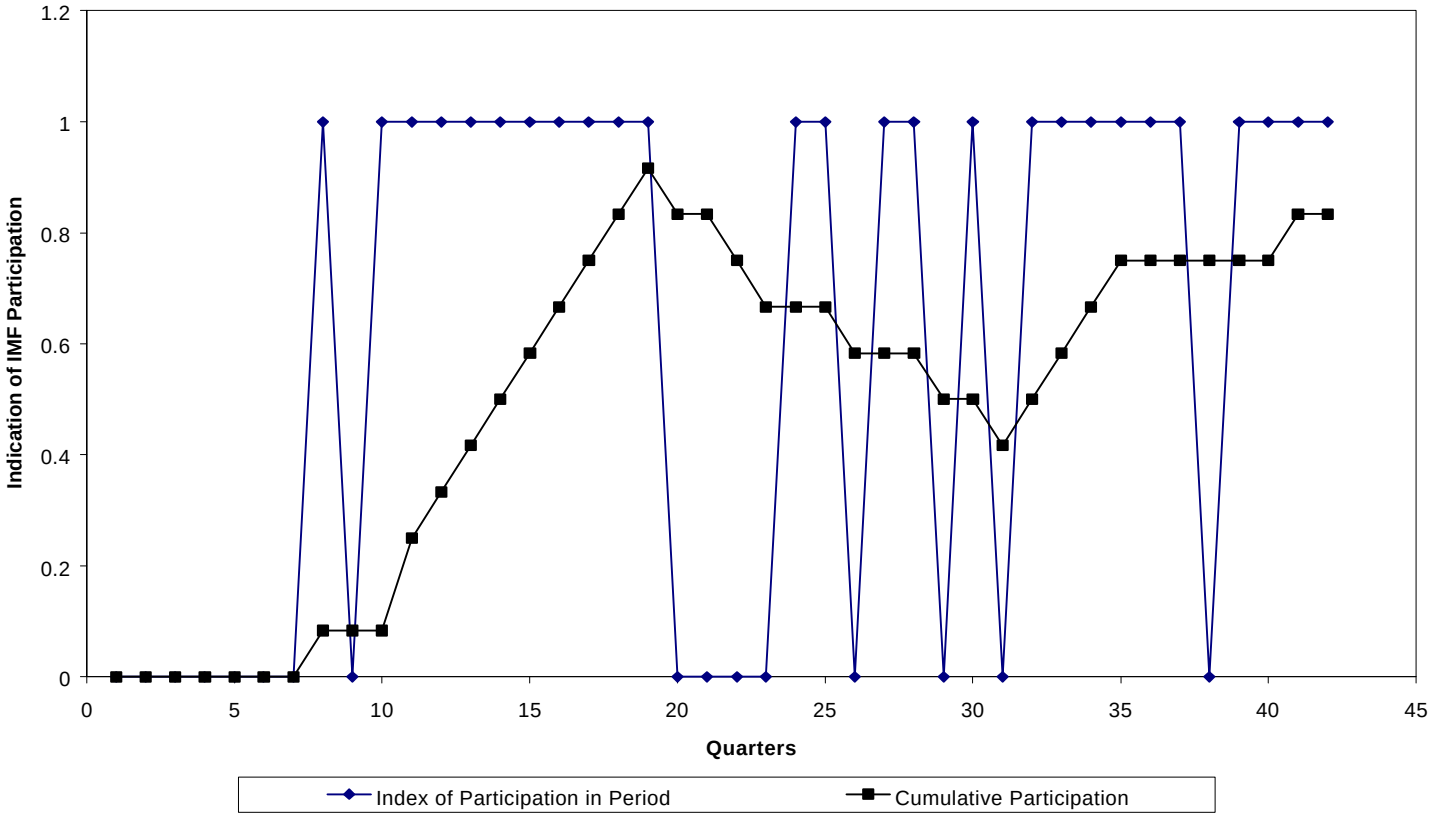

Table 1
Participation in IMF programs during sample period

Status of member relative to IMF	Number of observations	Percentage of total
No program participation in quarter:	4695	65.9
Program failure, no new program in quarter:	21	0.3
Program failure, adoption of a new program:	53	0.8
Participation in first tranche	252	3.5
Participation in stand-by arrangement	1251	17.6
Participation in EFF	342	4.8
Participation in SAF	186	2.6
Participation in ESAF	200	2.8
Participation in hybrid programs	121	1.7

Total observations: 7120.

Table 2
Transition Matrix in C_{Rit} Crisis Episodes

		period t	
		Crisis	No crisis
period t-1	Crisis	85.1 1696	14.9 296
	No crisis	7.9 303	92.1 3516

Total observations: 5811

Table 3
Correlation of Measures of Crisis

	I_{it}	C_{Rit}	C_{Git}	C_{Iit}
I_{it}	1.00000	0.23965	-0.08627	0.07564
		<.0001	<.0001	<.0001
	5349	5349	5349	5349
C_{Rit}	0.23965	1.00000	-0.07000	0.01112
	<.0001		0.0001	0.4162
	5349	5349	5349	5349
C_{Git}	-0.08627	-0.07000	1.00000	-0.01558
	<.0001	0.0001		0.2546
	5349	5349	5349	5349
C_{Iit}	0.07584	0.01112	-0.01558	1.00000
	<.0001	0.4162	0.2546	
	5349	5349	5349	

5349 observations, for 90 countries over 76 quarters.

Table 4
Determinants of the Conditional Probabilities p_{jxt} and p_{jNt} derived from Probit Estimation using the Reserve/Import Ratio measure of Crisis

	p_{ixt}			p_{iNt}	
	Simple	Fixed Effects		Simple	Fixed Effects
Intercept	-2.19 **		Intercept	0.94 **	
SE	0.18			0.27	
χ^2	143.54			11.95	
ρ_{j0}	0.64 **	0.63 **	ρ_{j0}	0.14	-0.36
SE	0.17	0.23		0.15	0.25
χ^2	14.72	7.38		0.93	2.19
I_{jt-1}	0.21	0.46 **	I_{jt-1}	-0.28	-0.40 *
SE	0.15	0.18		0.18	0.21
χ^2	1.99	7.01		2.53	3.61
IC_{jt-1}	-0.37 *	-0.32	IC_{jt-1}	0.67 **	0.65 **
SE	0.19	0.23		0.24	0.29
χ^2	3.59	1.90		7.77	4.92
R_{j0}	0.71 **	0.30	R_{j0}	-2.85 **	-3.18 **
SE	0.23	0.30		0.26	0.32
χ^2	9.36	1.00		119.27	96.42
m_{jxt-1}	0.55 **	0.17	m_{jNt-1}	1.00 **	0.18
SE	0.26	0.34	SE	0.24	0.31
χ^2	4.45	0.26	χ^2	17.82	0.36
Log Likelihood	-416.90	-346.66	Log Likelihood	-381.57	-332.61
χ^2 (58)		140.48 **	χ^2 (75)		97.92 **
Exit Number	166	166	Enter Number	158	158
Total Observed	1031	1031	Total Observed	2477	2477

Critical values of χ^2: 3.84 (95 percent level of confidence), significance marked by **
2.70 (90 percent level of confidence), significance marked by *

Fixed-effect terms were calculated in the appropriate estimations and are available on request. Given the presence of fixed-effect terms, no intercept was calculated.

Table 5 **Determinants of the Conditional Probability p_{jxt} derived from Proportional-Hazards Estimation: Reserve/Import Ratio Definition**

	White Correction	Stratified sample	Frailty measure
R_{i0}	1.21 **	-1.78 **	-0.35
SE	0.49	0.76	0.57
χ^2	6.15	5.42	0.37
Risk Ratio	3.38	0.17	0.70
ρ_{j0}	0.88 **	0.71	0.58
SE	0.25	0.62	0.41
χ^2	12.52	1.29	2.00
Risk Ratio	2.42	2.03	1.78
I_{it-1}	0.57 *	0.78 **	0.75 **
SE	0.31	0.37	0.30
χ^2	3.37	4.36	6.25
Risk Ratio	1.76	2.19	2.12
IC_{it-1}	-0.84 *	-0.00	-0.37
SE	0.45	0.48	0.41
χ^2	3.53	0.00	0.81
Risk Ratio	0.91	1.00	0.69
m_{jt-1}	0.67 **	0.01	0.46 *
SE	0.14	0.66	0.28
χ^2	22.61	0.00	2.70
θ			0.46 **
SE			0.06
χ^2 (θ=1)			94.90
Log Likelihood	-604.91	-208.97	-582.94
Exit Number	143	143	143
Total Observed Crisis	1087	1087	1087

Critical values of χ^2: 3.84 (95 percent level of confidence), significance marked by **
2.70 (90 percent level of confidence), significance marked by *
Standard errors in first column are corrected for unspecified heterogeneity with a White correction.

Table 6 **Determinants of the Conditional Probability p_{jNt} derived from Proportional-Hazards Estimation: Reserve/Import Ratio Definition**

	White Correction	Stratified sample	Frailty measure
R_{i0}	2.08 **	-1.44	-0.84
SE	0.82	1.11	0.85
χ^2	6.48	1.66	0.98
Risk Ratio	8.06	0.24	0.43
ρ_{j0}	1.58 **	3.08 **	1.68 **
SE	0.39	0.68	0.38
χ^2	16.30	20.85	19.54
Risk Ratio	4.88	21.84	5.38
I_{it-1}	-0.25	0.30	-0.01
SE	0.40	0.42	0.35
χ^2	0.14	0.48	0.00
Risk Ratio	0.86	1.34	0.99
IC_{it-1}	0.72	-0.36	0.18
SE	0.39	0.48	0.45
χ^2	2.06	0.48	0.16
Risk Ratio	1.06	1.34	1.14
m_{jt-1}	0.92 **	1.86 **	1.96 **
SE	0.14	0.58	0.47
χ^2	22.61	10.33	17.39
θ			0.69 **
SE			0.07
χ^2 (θ=1)			37.11
Log Likelihood	-589.02	-245.16	-574.40
Entry Number	123	123	123
Total NonCrisis	1087	1087	1087
Total Events	3721	3721	3721

Critical values of χ^2: 3.84 (95 percent level of confidence), significance marked by **
2.70 (90 percent level of confidence), significance marked by *

Standard errors in first column are corrected for unspecified heterogeneity with a White correction.

Appendix

A Reduced-form Model of the Evolution of the Reserve-Import Ratio

Foreign-exchange reserve holdings evolve according to the identity:

(1) $$R_t = R_{t-1} + m_t$$

with m_t representing the quantity of central-bank intervention to influence the foreign exchange market.

Imports (Z_t), by contrast, are assumed to evolve in a p^{th} order autoregressive process. The endogenous variable $h_t = (R_t/Z_t)$ is then a p^{th} order autoregression as well.

(2) $$h_t = \phi_1 h_{t-1} + \phi_2 h_{t-2} + \phi_3 h_{t-3} + ... + \phi_p h_{t-p} + m_t + \epsilon_t$$

Since the empirical record suggests use of a p=4 lag structure, I present the remaining derivations with that structure.

As Hamilton (1994, ch. 19) illustrates, this equation can be rewritten:

(3) $$\Delta h_t = -(\phi_2 + \phi_3 + \phi_4)\Delta h_{t-1} - (\phi_3 + \phi_4)\Delta h_{t-2} - \phi_4 \Delta h_{t-3} + (\phi_1 + \phi_2 + \phi_3 + \phi_4 - 1) h_{t-1} + m_t + \epsilon_t$$

indicating (for positive primitive parameters ϕ_i) negative coefficients on the lagged difference terms that decline in absolute value. If imports and reserves exhibit equal growth over time on average, then the summation $(\phi_1+\phi_2+\phi_3+\phi_4)$ will equal unity, leading to a zero coefficient on the lagged level.

These derivations were presented for exogenous m_t. In fact, even in the absence of crisis or IMF program participation, central banks will use foreign-exchange reserves to manipulate the foreign exchange market. The central bank's rule in such situations has been researched in some detail; Black (1985) provides a summary of conclusions. I introduce here a simple variant consistent with that research: that the central bank in country i uses its interventions to achieve a target reserve-import ratio h^*_i. Suppose that the central-bank loss function specified as the sum of the normalized cost of intervention and a quadratic in expected deviations from the target ratio.

(4) $$\Omega_t = m_t + \lambda(h_i^* - E_{t-1}h_t)^2$$

Through substitution, the central bank's intervention can be described as

(5) $$m_t = [(1/2\lambda) + h_i^*] - \phi_1 h_{t-1} - \phi_2 h_{t-2} - \phi_3 h_{t-3} - \phi_4 h_{t-4}$$

and the evolution of the reserve/import ratio will take the form

$$(6) \qquad \Delta h_t = [(1/2\lambda) + h_i^*] - (\phi_2 + \phi_3 + \phi_4)\Delta h_{t-1} - (\phi_3 + \phi_4)\,\Delta h_{t-2} - \phi_4\,\Delta h_{t-3} +$$

$$(\phi_2 + \phi_3 + \phi_4 - 1)\, h_{t-1} - \phi_2\, h_{t-2} - \phi_3\, h_{t-3} - \phi_4\, h_{t-4} + \epsilon_t$$

with country-specific intercept.

The existence of crisis will have two hypothesized effects on this economy. First, the nature of the autoregressive process in reserves (6) will be altered by the central bank's shortage of reserves. There will be import constraints and other trade restrictions that change the autoregressive process in imports; thus, the ϕ_i as estimated here will change. Second, the central bank will become more interventionist in leading to a return to desired reserve balances.

Participation in an IMF program also has two anticipated effects. First, the country's introduction of exchange-rate and expenditure reforms will increase the ϕ_i as reserves are accumulated at a more rapid pace. Second, the government's autonomy in manipulating reserves through a rule such as (5) will be greatly reduced.

The error-correction equation implied by these extensions is given in (7) for country i.

$$(7) \qquad \Delta h_{it} = \alpha_{io} + \Sigma_j\, (\beta_j + \gamma_{cj}\, C_{it-j} + \gamma_{Ij}\, D_{it-j})\, \Delta h_{it-j} + (\delta_o + \delta_c\, C_{it-1} + \delta_I\, D_{it-1})\, h_{it-1} + \epsilon_{it}$$

The basic error-correction structure is represented by the coefficients $\alpha_{io,}$ δ_o and β_j. α_{io} is the intercept vector, and represents country-specific differences in reserve accumulation, including differing central-bank preferences for target reserves across countries. β_j represents the autoregressive representation of response to lagged shocks, and is expected to be negative. δ_o is the error-correction coefficient, and its expected negative coefficient captures the gradual reversion of h_{it} toward its country-specific target value.

The specification of (7) includes as well the extent to which participating in IMF programs or being caught in crisis leads to a change in the behavioral response of the countries concerned. I define the variable D_{it} as an indicator function taking the value of unity when country i is in an IMF program in period t, and the variable C_{it} as an indicator function taking the value of unity when country i is in economic crisis in period t. Significant coefficients on terms in C_{it-j} and D_{it-j} will suggest that the behavioral forces (either market-based or in the central bank) driving the evolution of h_{it} are altered by participation in IMF program or by crisis. The existence of crisis will have two hypothesized effects. First, there will be import constraints and other trade restrictions that change the autoregressive process in imports; thus, the γ_{cj} are expected to be positive. Second, the central bank will become more interventionist in leading to a return to desired reserve balances, and δ_c will be negative.

Participation in an IMF program also has two anticipated effects. First, the country's introduction of exchange-rate and expenditure reforms will lead to γ_{Ij} taking the same sign as β_j as

reserves are accumulated at a more rapid pace. Second, the government's autonomy in manipulating reserves will be greatly reduced, leading to δ_I of opposite sign to δ_o. I investigate as well the possibility of a longer-term effect of participation in IMF programs in some analyses by including the variable p_{it} as a separate regressor. This variable measures the percentage of the previous three years that country i spent in IMF programs.

The data for each of the 86 countries were stacked for joint estimation, thus maintaining the hypothesis that the behavioral forces captured in this specification are common to all countries in the sample. The error ϵ_{it} is assumed normally and identically distributed. The initial estimation equation began with 3 lags of Δh_{it}, necessitating the exclusion of data from 1973.[23] In estimation including p_{it}, moreover, data from 1974 and 1975 were excluded as well. Explanatory variables were then pared by use of F tests of joint significance at the 95 percent confidence level. I report the "before" and "after" product of this estimating strategy in Table A3.[24] The second column of figures provides the coefficient estimates from the preferred specification.

The α_{io} coefficients (not reported) indicate significant differences in target reserve ratios across countries. All coefficient estimates were larger than zero, ranging from .52 (for Malta) to .003 (Sudan). Forty-six of the eighty-six country-specific coefficients were significantly different from zero, with significant differences across countries evident among many of the forty-six.[25] There is a significant autoregressive structure illustrated by the coefficients β_j of Δh_{it-j}, with the signs and magnitude as forecast by the model of the appendix. This coefficient is not significantly affected by immersion in a crisis period, as indicated by the coefficients on $C_{it-j}\Delta h_{it-j}$. However, those countries participating in IMF programs have a redoubled autoregressive response, as postulated above and as indicated by the coefficient γ_{I1} of the variable $D_{it-1}\Delta h_{it-1}$.

The error-correction coefficient is not significantly altered by the existence of crisis. This coefficient is, however, significantly reduced for countries participating in IMF programs, once again as postulated above, as indicated by the comparison of the coefficients δ_o and δ_I. There is a significant reversion effect captured in the error-correction coefficient δ_o of -0.13; its relatively

---

[23] The model derived in the appendix suggests the inclusion of four lags of the variable h_t as well. Given the arithmetic relationship between h_{it-1}, h_{it-2} and Δh_{it-2} (for example), such a specification is only identified through cross-coefficient parameter restrictions. When these restrictions (as found in appendix equation (A9)) were imposed, the restrictions were rejected by the data at the 95 percent level of significance. I thus fall back to the simpler form in equation (1) above.

[24] Country-specific intercepts are excluded from Table A3, but are available from the author on request.

[25] The variable p_{it} is not included in Table A3. When it is included it has a small negative coefficient, and a T statistic of .32 – insignificant at the 95 percent level of confidence. Its coefficient is larger, negative and significantly different from zero if the regression is run excluding the country-specific effects, suggesting that attention in future work need be paid to the determinants of those country-specific differences.

small size indicates partial quarterly adjustment to target levels. This speed of adjustment is, however, slowed further for those countries in IMF programs as the coefficient estimate $\delta_I = 0.04$ indicates. This suggests that the central-bank discretion in manipulating h_{it} is reduced for those countries participating in IMF programs.

This dynamic model of the evolution of h_{it} provides the starting point for a further investigation of the hypothesized interaction between IMF participation and crisis.

Table A1**: Sensitivity Analysis on Definition of Crisis**

The text reports the results from defining crisis with identical mean occurrence as participation in IMF programs. In the following tables I report the dynamics associated with different definitions. There are 4120 observations in each set.

Criterion: **33 percent of observations below critical level**.

		period t	
		Crisis	No crisis
period t-1	Crisis	83.2	16.8
		1136	229
	No crisis	8.3	91.7
		228	2527

Criterion: **20 percent of observations below critical level**.

		period t	
		Crisis	No crisis
period t-1	Crisis	81.4	18.6
		676	155
	No crisis	4.7	95.3
		154	3135

Criterion: **10 percent of observations below critical level**.

		period t	
		Crisis	No crisis
period t-1	Crisis	70.1	29.9
		291	124
	No crisis	3.3	96.7
		123	3582

Criterion: **5 percent of observations below critical level**.

		period t	
		Crisis	No crisis
period t-1	Crisis	63.9	36.1
		133	75
	No crisis	1.9	98.1
		74	3838

Criterion: **1 percent of observations below critical level**.

		period t	
		Crisis	No crisis
period t-1	Crisis	33.3	66.7
		14	28
	No crisis	0.7	99.3
		28	4050

Table A4.1
Determinants of the Conditional Probabilities p_{jxt} and p_{jNt} derived from Probit Estimation using the Reserve Growth measure

	p_{jxt}			p_{jNt}	
	Simple	Fixed Effects		Simple	Fixed Effects
Intercept	1.10 **		Intercept	-1.05 **	
SE	0.18			0.10	
χ^2	36.82			103.38	
ρ_{j0}	-0.15	-0.12	ρ_{j0}	0.04	0.10
SE	0.14	0.20		0.09	0.12
χ^2	1.13	0.34		0.21	0.74
I_{jt-1}	0.16	0.36 **	I_{jt-1}	-0.13	-0.20 **
SE	0.16	0.18		0.09	0.10
χ^2	1.06	3.97		2.00	3.85
IC_{jt-1}	-0.06	-0.14	IC_{jt-1}	0.16	0.14
SE	0.23	0.26		0.12	0.14
χ^2	0.07	0.28		1.63	1.07
R_{j0}	-0.30 **	-0.52 **	R_{j0}	0.24 **	0.22 **
SE	0.03	0.05		0.02	0.04
χ^2	100.75	94.06		108.88	41.12
m_{jxt-1}	0.19	0.59 **	m_{jNt-1}	0.24 **	0.01
SE	0.15	0.18	SE	0.06	0.07
χ^2	1.63	11.43	χ^2	13.30	0.01
Log Likelihood	-595.94	-528.34	Log Likelihood	-1517.80	-1451.51
χ^2 (57)		127.20	χ^2 (75)		132.58
Exit Number	788	788	Enter Number	776	776
Total Observed	1103	1103	Total Observed	2629	2629

Critical values of χ^2: 3.84 (95 percent level of confidence), significance marked by **
2.70 (90 percent level of confidence), significance marked by *

Fixed-effect terms were calculated in the appropriate estimations and are available on request. Given the presence of fixed-effect terms, no intercept was calculated.

Table A4.2

Determinants of the Conditional Probabilities p_{jxt} and p_{jNt} derived from Probit Estimation using the Inflation measure of crisis

	p_{jxt}			p_{jNt}	
	Simple	Fixed Effects		Simple	Fixed Effects
Intercept	-0.56 **		Intercept	-1.55 **	
SE	0.15			0.13	
χ^2	13.46			147.90	
ρ_{j0}	0.06	0.14	ρ_{j0}	0.09	-0.03
SE	0.12	0.16		0.12	0.18
χ^2	0.25	0.75		0.64	0.03
I_{jt-1}	0.11	0.12	I_{jt-1}	-0.03	-0.07
SE	0.12	0.16		0.11	0.13
χ^2	0.72	0.60		0.06	0.29
IC_{jt-1}	-0.10	-0.14	IC_{jt-1}	0.14	0.15
SE	0.17	0.22		0.16	0.18
χ^2	0.34	0.36		0.80	0.71
R_{j0}	0.11 **	0.13 **	R_{j0}	-0.00	-0.07
SE	0.03	0.05		0.03	0.04
χ^2	10.15	5.82		0.01	2.22
m_{jxt-1}	0.04	0.24	m_{jNt-1}	1.12 **	0.10
SE	0.17	0.22	SE	0.11	0.15
χ^2	0.06	1.19	χ^2	106.16	0.44
Log Likelihood	-684.94	-525.76	Log Likelihood	-1019.16	-847.68
χ^2 (75)		318.36 **	χ^2 (75)		344.04 **
Exit Number	385	385	Enter Number	377	377
Total Observed	1051	1051	Total Observed	2587	2587

Critical values of χ^2: 3.84 (95 percent level of confidence), significance marked by **
2.70 (90 percent level of confidence), significance marked by *

Fixed-effect terms were calculated in the appropriate estimations and are available on request. Given the presence of fixed-effect terms, no intercept was calculated.

Table A5.1: **Determinants of the Conditional Probability p_{ixt} derived from Proportional-Hazards Estimation: Reserve Growth Definition**

	White Correction	Stratified sample	Frailty measure
R_{i0}	-0.35 **	-0.47 **	-0.40
SE	0.07	0.06	0.57
χ^2	23.98	66.70	0.49
Risk Ratio	0.71	0.61	0.67
ρ_{j0}	-0.22 *	-0.37 *	0.56
SE	0.12	0.20	0.41
χ^2	3.35	3.39	1.86
Risk Ratio	0.81	0.62	1.75
I_{it-1}	0.29 **	0.58 **	0.75 **
SE	0.14	0.18	0.30
χ^2	4.04	9.59	6.25
Risk Ratio	1.38	1.78	2.11
IC_{it-1}	-0.13	-0.36	-0.37
SE	0.24	0.26	.41
χ^2	.55	1.66	0.81
Risk Ratio	0.92	0.90	0.69
m_{jxt-1}	-0.28	-0.05	0.07
SE	0.10	0.15	0.23
χ^2	8.47	0.10	0.09
θ			.99
SE			.01
χ^2 (θ=1)			2.78
Log Likelihood	-973.72	-625.32	-925.00
Exit Number	746	746	746
Total Obs. Crisis	1213	1213	1213

Critical values of χ^2: 3.84 (95 percent level of confidence), significance marked by **
2.70 (90 percent level of confidence), significance marked by *

Standard errors in first column are corrected for unspecified heterogeneity with a White correction.

Table A5.2: **Determinants of the Conditional Probability p_{ixt} derived from Proportional-Hazards Estimation: Inflation Crisis Definition**

	White Correction	Stratified sample	Frailty measure
R_{i0}	0.17 **	0.10	0.10 *
SE	0.06	0.07	0.06
χ^2	7.17	2.14	2.77
Risk Ratio	1.19	1.11	1.10
ρ_{j0}	0.16	-0.19	-0.17
SE	0.34	0.28	0.23
χ^2	0.19	0.43	0.54
Risk Ratio	1.16	0.83	0.84
I_{it-1}	0.18	-0.04	0.01
SE	0.30	0.25	0.20
χ^2	0.34	0.02	0.00
Risk Ratio	1.19	0.96	1.01
IC_{it-1}	-0.10	0.24	0.11
SE	0.30	0.36	0.28
χ^2	0.11	0.33	0.15
Risk Ratio	0.90	1.20	1.12
m_{jxt-1}	-1.22 **	-0.54	-0.66 **
SE	0.50	0.36	0.30
χ^2	5.95	2.23	4.84
θ			0.55 **
SE			0.04
χ^2 (θ=1)			180.35
Log Likelihood	-967.8	-420.38	-935.18
Exit Number	339	339	339
Total Crisis	960	960	960

Critical values of χ^2: 3.84 (95 percent level of confidence), significance marked by **
2.70 (90 percent level of confidence), significance marked by *

Standard errors in first column are corrected for unspecified heterogeneity with a White correction.

Table A6.1: **Determinants of the Conditional Probability p_{iNt} derived from Proportional-Hazards Estimation: Reserve Growth Definition**

	White Correction	Stratified sample	Frailty measure
R_{i0}	0.51 **	0.83 **	0.63 **
SE	0.09	0.08	0.05
χ^2	23.29	99.20	158.76
Risk Ratio	1.67	2.29	1.87
ρ_{j0}	0.34	0.32	0.15
SE	0.19	0.18	0.18
χ^2	3.12	3.24	0.69
Risk Ratio	1.40	1.38	1.16
I_{it-1}	-0.03	-0.07	-0.04
SE	0.14	0.17	0.15
χ^2	0.04	0.16	0.07
Risk Ratio	0.97	0.93	0.96
IC_{it-1}	0.14	-0.12	-0.02
SE	0.20	0.24	.13
χ^2	0.41	0.23	0.01
Risk Ratio	1.15	0.87	0.98
m_{jxt-1}	1.36	1.00	1.08
SE	0.10	0.14	0.13
χ^2	204.23	53.62	69.02
θ			.68
SE			.04
χ^2 (θ=1)			133.76
Log Likelihood	-2241.14	-1315.33	-2176.60
Enter Number	721	721	721
Total	2464	2464	2464

Critical values of χ^2: 3.84 (95 percent level of confidence), significance marked by **
2.70 (90 percent level of confidence), significance marked by *

Standard errors in first column are corrected for unspecified heterogeneity with a White correction.

Table A6.2: **Determinants of the Conditional Probability p_{iNt} derived from Proportional-Hazards Estimation: Inflation Crisis Definition**

	White Correction	Stratified sample	Frailty measure
R_{i0}	0.06	-0.02	-0.07
SE	0.13	0.11	0.09
χ^2	0.21	0.04	0.60
Risk Ratio	1.06	0.98	0.94
ρ_{j0}	0.76 **	-0.15	-0.01
SE	0.39	0.32	0.24
χ^2	3.85	0.23	0.00
Risk Ratio	2.15	0.86	0.98
I_{it-1}	0.44 **	0.77 **	0.61 **
SE	0.22	0.25	0.22
χ^2	3.88	9.15	7.69
Risk Ratio	1.56	2.16	1.84
IC_{it-1}	-0.38	-1.20	-0.80 **
SE	0.24	0.36	0.32
χ^2	1.84	10.48	6.25
Risk Ratio	0.77	0.30	0.45
m_{jxt-1}	2.47 **	1.76	1.88 **
SE	0.20	0.37	0.30
χ^2	153.43	22.81	39.27
θ			0.49 **
SE			0.04
χ^2 (θ=1)			282.73
Log Likelihood	-1241.83	-511.29	-1148.85
Exit Number	298	298	298
Total	2707	2707	2707

Critical values of χ^2: 3.84 (95 percent level of confidence), significance marked by **
2.70 (90 percent level of confidence), significance marked by *

Standard errors in first column are corrected for unspecified heterogeneity with a White correction.

Data Appendix

Sources of data used in this analysis

Data on reserves and imports are drawn from the CD-ROM compilation of the International Financial Statistics of the International Monetary Fund.

Series	Code
Reserves minus gold	1l.d
Gold holdings	1.and
Imports	71..d when available; otherwise, 71
Merchandise imports	78abd
Service imports	78aed
Income imports	78ahd
Exchange rate	rf

Data on IMF programs are drawn from the Annual Reports of the IMF. If a program were in place for more than 10 days of a quarter, then the country was considered participating in a program in that period.

Countries included in the data set are reported by availability for various regressions. (Some countries listed under hh_{it} or hg_{it} had a limited number of observations for those sensitivity analyses but were available in full sample for the h_{it} regression analysis.)

hh_{it}	hg_{it}		h_{it}	
Turkey	The preceding, plus		The preceding, plus	
South Africa	Benin *	Bahrain	Chad *	Tanzania
Argentina	Burkina Faso *	Cyprus	Congo *	Uganda
Bolivia	Cameroon *	Lebanon	Niger *	W. Samoa
Brazil	Central African Rep *	Egypt	Senegal *	Romania
Chile	Cote d'Ivoire *	Afghanistan	Liberia	Swaziland
Guatemala	Gabon *	Malaysia	Haiti	
Peru	Mali *	Algeria	Mexico	
Israel	Togo *	Burundi	Nicaragua	
Jordan	Malta	Ghana	Panama	
Bangladesh	Colombia	Kenya	Guyana	
Myanmar	Costa Rica	Malawi	Jamaica	
Sri Lanka	Dominican Rep.	Mauritania	Singapore	
India	Ecuador	Mauritius	Botswana	
Indonesia	El Salvador	Morocco	The Gambia	
Korea	Honduras	Nigeria	Lesotho	
Nepal	Paraguay	Tunisia	Madagascar	
Pakistan	Uruguay	Zambia	Rwanda	
Philippines	Venezuela	Fiji	Sierra Leone	
Thailand	Barbados	China	Somalia	
Ethiopia	Trinidad/Tobago		Sudan	

* - member of CFA franc zone