A. Data and Sample Construction

Tables A1, A2, and A3 present the data sources used to construct our panel of debt and price growth. Table A1 presents the equity price indices used to calculate 3-year real equity price growth.\(^1\) The equity data is drawn from four sources: Global Financial Data (GFD), the International Monetary Funds (IMF) International Financial Statistics, Bloomberg and the Jordá, Schularick and Taylor MacroHistory database (JST). Table A2 presents the house price indices used to calculate 3-year real house price growth. The data is drawn from Bank of International Settlements’ Property Price Statistics (BIS), the OECD’s Household Prices database (OECD), and JST. Finally, Table A3 presents the measures of debt outstanding used to calculate 3-year changes in debt to GDP for the non-financial business and household sector. The data is drawn from the International Monetary Funds’ Global Debt Database (IMF), the Bank of International Settlements’ Total Credit Statistics (BIS), and JST. Panel A presents an overview of the sources for business debt, while Panel B presents the sources for household debt.

When constructing our panel, we occasionally need to combine data from different sources to create time-series of changes within countries. When doing so, we calculate the changes within each data source to ensure smooth time-series. For example, to create a time-series of 3-year changes in real house prices for Denmark we combine data from the Jordá, Schularick and Taylor Macrohistory Database from 1950 to 1972, with data from the Bank of International Settlements’ Property Price Statistics from 1970 to 2018. Thus, in 1972 we calculate the 3-year change in real house prices for Denmark using only the JST data, and in 1973 we calculate the 3-year change in real house prices for Denmark using only the BIS data.

B. The Use of Linear Probability Models

The econometric specifications reported in the text are Linear Probability Models (LPM)—i.e., we run linear regressions where the dependent variable is a binary indicator.

\(^1\) We primarily rely on price indices to calculate equity price growth, but for a small subset of countries we use total return indices instead as no suitable price index was available. See Table A1 for details.
The use of LPMs offers a number of advantages relative to maximum likelihood models for handling binary outcomes such as Logit or Probit. First, the LPM coefficients are automatically interpretable as marginal effects, which show how the conditional probability of a future crisis changes as a function of the covariates. Second, LPMs allow us to include country fixed effects without encountering the incidental parameters problem. And, finally, LPM make it easier to conduct appropriate statistical inference on the marginal effects of interest (without using the delta method) by computing Driscoll-Kraay (1998) standard errors.

The downside of using LPMs is that they can generate predicted probabilities that lie outside the $[0, 1]$ interval. If the predicted probabilities fall outside the $[0, 1]$ interval for a significant portion of the sample, this can signal model misspecification, leading the LPM-implied marginal effects to diverge significantly from those implied by Logit and Probit models.

Fortunately, it turns out that the use of Linear Probability Models (LPMs) makes almost no difference in our setting: We obtain nearly identical marginal effects using Logit or Probit models.

Let us now explain in greater detail.

**Analysis without country fixed effects:** If we omit the country fixed-effects, Logit and Probit models deliver the exact same marginal effects as LPMs in our setting.

To see why, note that the covariates in our specifications are a set of binary indicators that define a discrete partition of the sample. As a result, the predicted probabilities from our LPMs are just the empirical historical probabilities in each cell of the partition and, thus, always lie in the $[0,1]$ interval. Our estimated LPM coefficients are differences between these empirical conditional probabilities and, thus, are proper marginal effects.

Going further, since the covariates define a discrete partition of the sample, any maximum likelihood model for handling binary outcome variables—including Logit or Probit—is going to yield the same predicted probabilities as a LPM and hence the same marginal effects.

Formally, our baseline regressions divide country-years into the following 2-by-2 partition:

<table>
<thead>
<tr>
<th>Debt growth</th>
<th>Low</th>
<th>High</th>
</tr>
</thead>
<tbody>
<tr>
<td>Price growth</td>
<td>Low</td>
<td>$\hat{p}<em>{LL} = n</em>{L,L}/N_{L,L}$</td>
</tr>
<tr>
<td>High</td>
<td>$\hat{p}<em>{H,L} = n</em>{H,L}/N_{H,L}$</td>
<td>$\hat{p}<em>{H,H} = n</em>{H,H}/N_{H,H}$</td>
</tr>
</tbody>
</table>

where $N$ is the number of country-years in each cell in our sample, $n$ is the number of country-years in each cell in which a crisis later materializes, and $\hat{p}$ is the in-sample conditional probability of a crisis in that cell. For instance, $\hat{p}_{H,H}$ is the sample probability that a crisis arrives within the next 3 years conditional on starting from inside the $R$-zone.

If we then estimate the following multivariate LPM model:

\[
\text{Crisis}_{t+1}^{t+3} = \alpha + \beta \cdot \text{High-Debt-Growth}_{lt} + \delta \cdot \text{High-Price-Growth}_{lt} + \gamma \cdot \text{R-zone}_{lt} + \epsilon_{t+1}^{t+3},
\]

the coefficient estimates are $\hat{\alpha} = \hat{p}_{L,L}$, $\hat{\beta} = \hat{p}_{L,H} - \hat{p}_{L,L}$, $\hat{\delta} = \hat{p}_{H,L} - \hat{p}_{L,L}$, and $\hat{\gamma} = (\hat{p}_{H,H} - \hat{p}_{L,H}) - (\hat{p}_{H,L} - \hat{p}_{L,L})$. As noted, the estimated coefficients from our LPMs are just differences—or differences-in-differences—between empirical conditional probabilities and, thus,
are proper marginal effects. Similarly, if we consider a Logit or Probit model in which

\[
\Pr(Crisis_{i,t+1} = \alpha + \gamma \cdot \text{R-zone}_t + e_{i,t+1} \mid \text{High-Debt-Growth}_t + \text{High-Price-Growth}_t, \text{R-zone}_t, \text{High-Debt-Growth}_t + \text{High-Price-Growth}_t)
\]

this yields the exact same estimated marginal effects—i.e., \( \hat{\alpha} = \hat{\beta}_{LL}, \hat{\beta} = \hat{\beta}_{LH} - \hat{\beta}_{LL}, \hat{\delta} = \hat{\beta}_{HL} - \hat{\beta}_{LL}, \) and \( \hat{\gamma} = (\hat{\beta}_{HH} - \hat{\beta}_{LH}) - (\hat{\beta}_{HL} - \hat{\beta}_{LL}). \)

Adding country fixed effects: Next, we consider what happens once we add country fixed effects. As we emphasize on page 12 of the paper, the inclusion of country fixed effects in our LPMs has almost no impact on the coefficients of interest—i.e., on the LPM-implied marginal parameters, which are of less economic interest. Specifically, letting \( i \)

\[
\text{Crisis}_{i,t+1} = \alpha + \gamma \cdot \text{R-zone}_t + e_{i,t+1} \mid \text{High-Debt-Growth}_t + \text{High-Price-Growth}_t, \text{R-zone}_t, \text{High-Debt-Growth}_t + \text{High-Price-Growth}_t
\]

we obtain \( \hat{\gamma} = (\hat{\beta}_{HH} - \hat{\beta}_{LH}) \) where \( \hat{\beta}_{HH} = (n_{LH} + n_{HL} + n_{HH})/(n_{LL} + n_{LH} + n_{HH}). \)

Of course, these estimated marginal effects are different from the estimates of the underlying Logit and Probit parameters, which are of less economic interest. Specifically, letting \( \Lambda[z] = \exp[z]/(1 + \exp[z]) \) denote the logistic function, the Logit parameters satisfy:

\[
\Pr(Crisis_{i,t+1} = \alpha + b \cdot \text{High-Debt-Growth}_t + d \cdot \text{High-Price-Growth}_t + g \cdot \text{R-zone}_t).
\]

Thus, the discrete marginal effects for the Logit model are given by the following expressions:

\[
\hat{\alpha} = \Lambda[\hat{\alpha}] = \hat{\beta}_{LL},
\hat{\beta} = \Lambda[\hat{\alpha} + \hat{\beta}] - \Lambda[\hat{\alpha}] = \hat{\beta}_{HL} - \hat{\beta}_{LL},
\hat{\delta} = \Lambda[\hat{\alpha} + \hat{\delta}] - \Lambda[\hat{\alpha}] = \hat{\beta}_{HL} - \hat{\beta}_{LL},
\hat{\gamma} = (\Lambda[\hat{\alpha} + \hat{\beta} + \hat{\delta}] - \Lambda[\hat{\alpha} + \hat{\beta}]) - (\Lambda[\hat{\alpha} + \hat{\delta}]) = (\hat{\beta}_{HH} - \hat{\beta}_{LH}) - (\hat{\beta}_{HL} - \hat{\beta}_{LL}).
\]

Similarly, the underlying Probit parameters satisfy:

\[
\Pr(Crisis_{i,t+1} = \Phi[a + b \cdot \text{High-Debt-Growth}_t + d \cdot \text{High-Price-Growth}_t + g \cdot \text{R-zone}_t],
\]

where \( \Phi[z] \) is the standard normal CDF. And, the marginal effects take the prior form replacing \( \Phi[z] \) with \( \Lambda[z] \).

4 The fixed-effect Logit model takes the form

\[
\Pr(Crisis_{i,t+1} = \alpha_i + b \cdot \text{High-Debt-Growth}_t + d \cdot \text{High-Price-Growth}_t + g \cdot \text{R-zone}_t),
\]

where \( \alpha_i \) is the fixed effect for country \( i \). Due to the inclusion of country fixed effects, the marginal effects vary across countries. Specifically, the marginal effects for country \( i \) are:

\[
\hat{\beta}_i = \Lambda[\hat{\alpha}_i + \hat{\beta}_i] - \Lambda[\hat{\alpha}_i],
\hat{\delta}_i = \Lambda[\hat{\alpha}_i + \hat{\delta}_i] - \Lambda[\hat{\alpha}_i],
\hat{\gamma}_i = (\Lambda[\hat{\alpha}_i + \hat{\beta}_i + \hat{\delta}_i] - \Lambda[\hat{\alpha}_i + \hat{\beta}_i]) - (\Lambda[\hat{\alpha}_i + \hat{\delta}_i] - \Lambda[\hat{\alpha}_i]).
\]

We obtain the average Logit-implied marginal effects by computing the average of these country-specific marginal effects across all country-years in our sample. (Thus, countries with a longer time-series receive a
Table A4: Estimated Marginal Effects: LPM versus Logit or Probit Models

<table>
<thead>
<tr>
<th>Model</th>
<th>Fixed effects</th>
<th>Countries</th>
<th>N</th>
<th>High credit</th>
<th>Multivariate</th>
<th>R-zone</th>
<th>Univariate</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td></td>
<td></td>
<td></td>
<td>(t)</td>
<td>High price</td>
<td>R-zone</td>
<td></td>
</tr>
<tr>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Panel A: Business sector</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>(1)</td>
<td>LPM</td>
<td>No</td>
<td>42</td>
<td>1,258</td>
<td>13.44</td>
<td>6.91</td>
<td>(2.67)***</td>
</tr>
<tr>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td>(1.11)</td>
<td>19.11</td>
<td>(2.66)***</td>
</tr>
<tr>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td>35.44</td>
<td>(3.46)***</td>
</tr>
<tr>
<td>(2)</td>
<td>Logit</td>
<td>No</td>
<td>42</td>
<td>1,258</td>
<td>13.44</td>
<td>6.91</td>
<td>(2.67)***</td>
</tr>
<tr>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td>(1.11)</td>
<td>19.11</td>
<td></td>
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<td></td>
<td></td>
<td></td>
<td>35.44</td>
<td>(3.46)***</td>
</tr>
<tr>
<td>(3)</td>
<td>Probit</td>
<td>No</td>
<td>42</td>
<td>1,258</td>
<td>13.44</td>
<td>6.91</td>
<td>(2.67)***</td>
</tr>
<tr>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td>(1.11)</td>
<td>19.11</td>
<td></td>
</tr>
<tr>
<td></td>
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<td></td>
<td></td>
<td></td>
<td>35.44</td>
<td>(3.46)***</td>
</tr>
<tr>
<td>(4)</td>
<td>LPM</td>
<td>Yes</td>
<td>42</td>
<td>1,258</td>
<td>11.47</td>
<td>6.91</td>
<td>(1.09)</td>
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<td>(1.10)</td>
<td>19.11</td>
<td>(2.66)***</td>
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<td></td>
<td></td>
<td>35.44</td>
<td>(3.46)***</td>
</tr>
<tr>
<td>(5)</td>
<td>LPM*</td>
<td>Yes</td>
<td>29</td>
<td>1,103</td>
<td>12.80</td>
<td>7.37</td>
<td>(1.04)</td>
</tr>
<tr>
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<td></td>
<td>(1.04)</td>
<td>19.75</td>
<td>(2.86)***</td>
</tr>
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<td></td>
<td>35.80</td>
<td>(3.34)***</td>
</tr>
<tr>
<td>(6)</td>
<td>Logit*</td>
<td>Yes</td>
<td>29</td>
<td>1,103</td>
<td>12.79</td>
<td>8.59</td>
<td>(1.14)</td>
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<td></td>
<td>(1.14)</td>
<td>20.09</td>
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<td></td>
<td></td>
<td>35.28</td>
<td>(3.34)***</td>
</tr>
<tr>
<td>(7)</td>
<td>Probit*</td>
<td>Yes</td>
<td>29</td>
<td>1,103</td>
<td>13.57</td>
<td>8.99</td>
<td>(1.19)</td>
</tr>
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<td></td>
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<td></td>
<td></td>
<td></td>
<td>(1.19)</td>
<td>18.97</td>
<td>(1.19)</td>
</tr>
</tbody>
</table>

| Panel B: Household sector |
| (1)   | LPM           | No        | 40 | 1,107       | 11.70        | -0.47  | (2.01)**   |
|       |               |           |    |             | (0.17)       | 17.69  | (2.45)**   |
|       |               |           |    |             |              | 27.78  | (3.03)**   |
| (2)   | Logit         | No        | 40 | 1,107       | 11.70        | -0.47  | (2.01)**   |
|       |               |           |    |             | (0.17)       | 17.69  | (2.45)**   |
|       |               |           |    |             |              | 27.78  | (3.03)**   |
| (3)   | Probit        | No        | 40 | 1,107       | 11.70        | -0.47  | (2.01)**   |
|       |               |           |    |             | (0.17)       | 17.69  | (2.45)**   |
|       |               |           |    |             |              | 27.78  | (3.03)**   |
| (4)   | LPM           | Yes       | 40 | 1,107       | 9.13         | 0.00   | (2.27)**   |
|       |               |           |    |             | (0.00)       | 20.94  | (3.18)**   |
|       |               |           |    |             |              | 28.60  | (3.44)**   |
| (5)   | LPM*          | Yes       | 25 | 991         | 9.48         | 0.16   | (2.22)**   |
|       |               |           |    |             | (0.04)       | 22.68  | (3.22)**   |
|       |               |           |    |             |              | 30.85  | (3.62)**   |
| (6)   | Logit*        | Yes       | 25 | 991         | 9.73         | 0.08   | (2.22)**   |
|       |               |           |    |             | (0.04)       | 23.19  | (3.22)**   |
|       |               |           |    |             |              | 31.20  | (3.62)**   |
| (7)   | Probit*       | Yes       | 25 | 991         | 8.83         | 0.13   | (2.22)**   |
|       |               |           |    |             | (0.04)       | 23.18  | (3.22)**   |
|       |               |           |    |             |              | 30.45  | (3.62)**   |

* These rows drop all countries where there is no variation in the binary outcome—i.e., countries that never experienced a crisis. Fixed-effects MLEs can perfectly predict the outcomes in these countries by setting \( \hat{\alpha} \to -\infty \), so these countries play no role in identifying the shared coefficients and marginal effects from fixed-effects MLEs.

Table A4 compares the marginal effects implied by LPM, Logit, and Probit models. We are predicting whether a financial crisis will arrive in the next 3 years and the format follows Table 5. Consider the results for the business sector in Panel A. The first three rows compare LPM, Logit, and Probit marginal effects when we omit the country fixed effects. As noted above, these are identical by necessity. Row (4) show the LPM-implied marginal effects once we include country fixed effects. Comparing the results in row (4) to those in rows (1), we see that the inclusion of country fixed effects has almost no impact on our LPM estimates. In row (5), we estimate an LPM after dropping the 13 countries in our sample that never experience a crisis: these countries will be dropped in fixed-effect Logit and Probit models. This has almost no impact. Finally, row (6) shows that the fixed-effects Logit yields very similar marginal effects as the fixed-effects LPM in row (5). And, row (7) shows that the fixed-effects Probit also generates nearly identical marginal effects as the fixed-effects LPM.

C. Cumulative and Incremental Probabilities of Crisis Onset at Different Horizons

In our setting, one can roughly infer the incremental probability of crisis onset at different horizons by tracking how the cumulative probability of crisis onset grows with the forecast horizon. Specifically, since it is rare to have multiple distinct crises in the same country over a short period of time, for small \( h \) we have:

\[
\text{Crisis}_{t+1 \text{ to } t+h} \equiv \max \{ \text{Crisis-Start}_{t+1, \ldots, t+h} \} \\
\approx \sum_{k=1}^{h} \text{Crisis-Start}_{t+k}.
\]

Equivalently, this means that:

larger weight in this average). We obtain Probit-implied marginal effects in an analogous fashion.
The first three years for an event. However, this is not statistically different from the probability of onset in the first or second year using the BVX crisis indicator when \(D \in \text{elevated} \) for three years following a business
Panels A2 and A3 both show that the incremental probability of crisis onset remains forecast horizon until year three. And, as expected, the coefficients in the Panels A2 and A3 that the cumulative probability of crisis arrival in Panel A climbs steadily with the \(p\)-values and has better small-sample properties than traditional Gaussian asymptotic theory. Kiefer and Vogelsang (2005) show that their fixed-
As shown in Panel B, a similar pattern holds for the household sector.
Probabilities—by comparing the coefficients from regressions involving \(Crisis_{t+1}^{i} \) and \(Crisis_{t+1}^{i-1} \) across the columns in our existing tables. (These statements are exact using the BVX crisis indicator when \(h = 2\), but only approximate when \(h = 3\) and 4.)
Panels A2 and A3 both show that the incremental probability of crisis onset remains elevated for three years following a business \(R\)-zone event. This corresponds to the fact that the cumulative probability of crisis arrival in Panel A1 climbs steadily with the forecast horizon until year three. And, as expected, the coefficients in the Panels A2 and A3 are very similar.

As shown in Panel B, a similar pattern holds for the household sector.

D. Using the Bootstrap to Assess the Impact of Small-Sample Statistical Problems

When estimating \(h\)-year forecasting regressions for \(h > 1\), we use Driscoll-Kraay (1998) standard errors allowing for residual correlation at up to \(m = \text{ceiling}(1.5 \times h)\) lags. We compute \(p\)-values using the “fixed-\(b\)” asymptotic theory of Kiefer and Vogelsang (2005). Fixed-\(b\) asymptotics were designed to address the tendency for traditional statistical tests based on heteroskedasticity and autocorrelation consistent (HAC) standard errors to over-reject in finite samples. Intuitively, HAC variance estimators tend to be quite noisy in small samples because of the need to non-parametrically estimate a series of higher order autocorrelations. And, the resulting noise in the variance estimator leads \(t\)-statistics based on HAC standard errors to have fatter-than-Gaussian tails in small samples. Kiefer and Vogelsang (2005) show that their fixed-\(b\) asymptotic theory delivers more conservative \(p\)-values and has better small-sample properties than traditional Gaussian asymptotic theory.

---

5 For the business \(R\)-zone, the marginal probability of onset is highest in the third year following an \(R\)-zone event. However, this is not statistically different from the probability of onset in the first or second year after an \(R\)-zone event. For the household, \(R\)-zone the marginal probability of onset is roughly constant for the first three years for an \(R\)-zone event.
6 If anything, since \(Crisis-Start_{t+h}^{i} \geq Crisis_{t+1}^{i} - Crisis_{t+1}^{i-1} \), the coefficients in Panel C are slightly larger than those in Panel B for \(h = 3\) and 4.
7 In contrast to the traditional asymptotic theory for inference for HAC inference, which is derived under the assumption that \(m \to \infty \) and \(m/T \to 0\) as \(T \to \infty\), Kiefer and Vogelsang (2005) assume that \(m = bT\) for some \(b \in (0,1)\)—i.e., that bandwidth \(m\) is a fixed fraction \(b\) of the length \(T\) of the panel. Let \(t_{b} = (\hat{\theta} - \theta_{0})/\sqrt{\hat{V}}\) denote the \(t\)-statistic using the Driscoll-Kraay (1998) variance estimator with bandwidth \(m = bT\). Then as \(T \to \infty\), Kiefer and Vogelsang (2005) show that \(t_{b} \to b(1)/\sqrt{Q(b)}\), where \(B(r)\) is a standard Brownian motion, \(\hat{B}(r) = B(r) - rB(1)\), and \(Q(b) = (2/b) \int_{0}^{1} B(r)dr - (2/b) \int_{0}^{1} \hat{B}(r + b)B(r)dr\). Furthermore, they show that \(Q(b) \to 1\) as \(b \to 0\), so that these
Gonglaves and Vogelsang (2008) show that inference using this fixed-\(b\) approach is asymptotically equivalent to inference using a moving-block bootstrap. However, for smaller samples, Gonglaves and Vogelsang (2008) argue that better approximations can be obtained via a moving-block bootstrap with a suitably chosen block length. Thus, to better assess statistical significance, we use a nonparametric bootstrap to estimate the small-sample distribution of our \(t\)-statistics.

Formally, we create a large number of bootstrap samples by resampling observations from our original dataset. We then estimate the unknown finite-sample distribution of \(t = (\hat{\theta} - \theta_0)/\sqrt{V}\) using the distribution of \(t^* = (\hat{\theta}^* - \hat{\theta})/\sqrt{\hat{V}}\) across these bootstrap samples, where \(\hat{\theta}^*\) and \(\hat{V}\) are the parameter estimate and its variance estimate in a given bootstrap sample. A major benefit of this bootstrap-\(t\) procedure (Efron (1982) and Hall (1988)) is that it allows us to assess the impact of a host of small-sample statistical problems in one fell swoop. These problems include (1) any small-sample biases in our coefficient estimates due Stambaugh (1999) bias or the fact that our predictors make use of full-sample quantiles; (2) the fact that sampling variation in the variance estimates leads the distribution of \(t\)-statistics to have fat tails in small samples; and (3) any correlation between coefficient and variance estimates which can lead \(t\)-statistics to have non-symmetric distributions in small samples.

For each bootstrap iteration, we construct a pseudo country-year panel dataset using a panel moving-blocks procedure (detailed below) that resamples blocks of temporally contiguous cross-sections from our original panel. For our 3-year crisis forecasting regressions the cross-section at time \(t\) is \(Z_t\). Letting \(\Delta_3\) (Debt/GDP)\(_{it}\) and \(\Delta_3\)Log(Price\(_{it}\))\(_{i}\in S_t\) where \(S_t\) is the set of countries in the panel at time \(t\). After constructing this pseudo panel, we redefine our \(High-Debt-Growth\), \(High-Price-Growth\), and \(R-zone\) indicator variables based on the quantiles in this pseudo panel. We then estimate our forecasting regression and compute Driscoll-Kraay (1998) standard errors using this pseudo panel dataset and save the resulting bootstrapped \(t\)-statistics: \(t^* = (\hat{\theta}^* - \hat{\theta})/\sqrt{\hat{V}}\). We compute bootstrap-implied p-values by comparing the \(t\)-statistic we obtain from our actual country-year panel to the resulting distribution of bootstrapped \(t\)-statistics.

To preserve the cross-sectional and time-series dependence within our country-year panel dataset, we create pseudo panel datasets by adapting the stationary moving blocks bootstrap of Politis and Romano (1994). We adapt the standard procedure slightly to deal with the unbalanced nature of our panel—i.e., the fact that number of observations in each cross-section, \(N_t\), varies over time.

Let \(Z_t\) for \(t = 1, \ldots, T\) denote the \(T\) cross-sections in our original panel. Let \(B_{t,k} = \{Z_t, Z_{t+1}, \ldots, Z_{t+k-1}\}\) be the block of \(k\) consecutive cross-sections starting from time \(t\). If \(t + i > T\) for some \(i \leq k - 1\), we let \(Z_{t+i} = Z_{\text{mod}(t+i,T)}\)—i.e. we “wrap the data around the circle”. Letting \(\{L_i\}\) be a sequence of iid draws from the geometric distribution with probability \(q\) and \(\{L_i\}\) be a sequence of iid draws from the discrete uniform distribution on \(\{1, 2, \ldots, T\}\), we create a pseudo panel by resampling blocks of \(random\) length as \(\{B_{t_1,L_1}, B_{t_2,L_2}, \ldots\}\). This process is stopped once \(T\) cross-sections have been selected. As discussed in Politis and Roman (1994), this procedure ensures that all cross-sections are equally likely to be resampled. This procedure also ensures that the pseudo panels will be stationary if the original panel is stationary.

We adapt this standard moving-blocks procedure to deal with the unbalance nature of our panel. The expected number of country-years from these pseudo equals the number of country years in our original panel. However, while the time-series length of these panels is fixed at \(T\), the number fixed \(b\) asymptotics converge to standard asymptotics in the limit where \(b \to 0\).
of countries years be distributed in a roughly bell-curved shape about this mean. To ensure that our pseudo panels have a similar number of country years, we only keep those pseudo panels where the number of countries years is within plus or minus one standard deviation (across bootstrap replications) of the mean (across bootstrap replications). Since the width of the cross-sections, \( N_t \), is roughly linear in time \( t \), this has an almost negligible impact on the likelihood that a given cross-section ends up in one of our pseudo-panels. However, this means that we throw out both pseudo panels that over-represent the earlier years in our sample when \( N_t \) was smaller as well those that over-represent the later years when \( N_t \) is larger.

Let \( D^* = (\hat{\theta} - \theta) / \sqrt{V} \) and let \( T_{\text{Boot}}(\alpha) \) denote the \( \alpha \) percentile of this distribution—i.e., \( \Pr[t^* \leq T_{\text{Boot}}(\alpha)] = \alpha \). Consider an asymmetric 2-tailed confidence interval with coverage probability \( 1 - \alpha \). We have

\[
1 - \alpha = \Pr[T_{\text{Boot}}(\alpha/2) \leq (\hat{\theta} - \hat{\theta})/\sqrt{V} \leq T_{\text{Boot}}(1 - \alpha/2)]
\approx \Pr[T_{\text{Boot}}(\alpha/2) \leq (\hat{\theta} - \theta_0)/\sqrt{V} \leq T_{\text{Boot}}(1 - \alpha/2)]
= \Pr[\hat{\theta} - \sqrt{V} \times T_{\text{Boot}}(1 - \alpha/2) \leq \theta_0 \leq \hat{\theta} - \sqrt{V} \times T_{\text{Boot}}(\alpha/2)].
\]

Thus, the implied bootstrapped \( p \)-value for a test of the hypothesis that \( \theta_0 = 0 \) is given by

\[
\alpha_{\text{Boot}} = \min\{2 \times T_{\text{Boot}}^{-1}(\hat{\theta}/\sqrt{V}), 2 \times (1 - T_{\text{Boot}}^{-1}(\hat{\theta}/\sqrt{V}))\}.
\]

Since \( t \)-statistics using consistent and asymptotically normal estimators are “asymptotically pivotal”—they are asymptotically standard normal regardless of the true underlying parameters—using these bootstrapped \( p \)-values offers asymptotical refinement in finite samples.

These results are presented in Table A6. We use 100,000 replications for each regression and a parameter of \( q = 1/8 \), so that the average block length is 8 years. Similar results obtain for other choices of \( q \). While the \( p \)-values derived from the bootstrap-\( t \) procedure are larger (i.e., less significant) than those based on asymptotic theory, we find that 3-year forecasting results are significant at the 5% level or better. Thus, \( t \)-statistics as large as those shown in Table A6 are highly unlikely to obtain by chance, even in finite samples.

We also use this bootstrapping procedure to assess any finite-sample bias of our coefficient estimates, whether due Stambaugh (1999) bias or the fact that our predictors make use of full-sample quantiles. Specifically, we report the standard bootstrap bias estimator \( \text{bias}(\hat{\theta}) = E_{\text{Boot}}[\hat{\theta}^* - \hat{\theta}] \) and the bias-adjusted estimate of \( \theta_0 \), \( \hat{\theta}_{\text{adjust}} = \hat{\theta} - \text{bias}(\hat{\theta}) = 2\hat{\theta} - E_{\text{Boot}}[\hat{\theta}^*] \). As shown in Table A6, our bootstrapping exercise suggests that the magnitude of these estimation biases is negligible.

E. Varying the Cutoffs Used to Define High Debt Growth and High Price Growth

To address concerns about functional-form overfitting, Table A7 asks whether our results are sensitive to the cutoffs we use to construct our indicators for high debt growth and high price growth. Table A7 shows that there is nothing special about the particular cutoffs we use to construct these indicators: we obtain similar results in the full sample, the pre-2000 sample, and the post-2000 sample for a variety of different cutoff values.

For various cutoff pairs \((c_D, c_P)\), we recompute:

\[
\text{High-Debt-Growth}_{it}(c_D) = 1\{\Delta_3(\text{Debt/GDP})_it > c_D\}
\]
\[
\text{High-Price-Growth}_{it}(c_P) = 1\{\Delta_3\log(\text{Price}_{it}) > c_P\}
\]
\[ R\text{-zone}_{it}(c_D,c_P) = 1\{\Delta_3(\text{Debt/GDP})_{it} > c_D\} \times 1\{\Delta_3\log(\text{Price}_{it}) > c_P\}. \]

Using the resulting indicator variables, we then rerun our baseline forecasting regressions in different samples:

\[
\text{Crisis}_{i,t+1 \text{ to } t+3} = \alpha_i + \beta \cdot \text{High-Debt-Growth}_{it}(c_D) \\
+ \delta \cdot \text{High-Price-Growth}_{it}(c_P) + \gamma \cdot R\text{-zone}_{it}(c_D,c_P) + \varepsilon_{i,t+1 \text{ to } t+3}
\]

We consider the full sample, the pre-2000 subsample (the last 3-year forecast is in 1996), and the post-2000 subsample (the first forecast is in 1997). For each sample, we report the estimated coefficient \( \hat{\gamma} \) on \( R\text{-zone}_{it}(c_D,c_P) \) as a function of the cutoffs \( (c_D,c_P) \) from a univariate regression and a multivariate regression that controls for \( \text{High-Debt-Growth}_{it}(c_D) \) and \( \text{High-Price-Growth}_{it}(c_P) \). The highlighted cells correspond to our baseline variable definitions. Panel A shows the results for the business sector and Panel B shows those for the household sector.

Table A7 shows that our central conclusion—the combination of rapid credit growth and asset price growth is associated with an elevated risk of a financial crisis—holds up in all three samples for a variety of different cutoff pairs \( (c_D,c_P) \). As one would expect based on the nonlinear relationship between past debt and asset price growth and the probability of a future crisis that we emphasize throughout, the estimated coefficient \( \hat{\gamma} \) on \( R\text{-zone}_{it}(c_D,c_P) \) and the associated \( t \)-statistic are generally increasing in both the cutoff for debt growth \( (c_D) \) and the cutoff for asset price growth \( (c_P) \). Compared to other cutoff values, our baseline definitions yield neither the strongest nor the weakest results. To be sure, there are specifications where \( \hat{\gamma} \) is no longer statistically significant in some subsamples, but our reading of Table A7 is that our key findings are highly robust to the particular choice of cutoffs.

### F. Robustness Checks for Table 7

Here we perform several robustness checks on Table 7 in the main text. Specifically, in Table 7, we estimated regressions of the form:

\[
\text{Crisis}_{i,t+1 \text{ to } t+h} = \alpha_i^{(h)} + \gamma^{Bus(h)} \cdot \text{Local R\text{-zone}}^{Bus}_{it} + \xi^{Bus(h)} \cdot \text{Global R\text{-zone}}^{Bus}_{it} \\
+ \gamma^{HH(h)} \cdot \text{Local R\text{-zone}}^{HH}_{it} + \xi^{HH(h)} \cdot \text{Global R\text{-zone}}^{HH}_{it} + \varepsilon_{i,t+1 \text{ to } t+h}
\]

for \( h = 1, 2, 3, \) and 4. In the results reported in Table 7, the \textit{Global R\text{-zone}} variable is simply the equal-weighted average across all countries in each annual cross-section. However, our results are quite robust to the way that we construct our \textit{Global R\text{-zone}} variable. First, as shown in Table A8 below, the results in Table 7 are very similar if the \textit{Global R\text{-zone}} variable for each country-year is defined as the equal-weighted fraction of \textit{other} sample countries that are in the \textit{R\text{-zone}} in that year—i.e., in a “leave one out” fashion. Second, as shown in Table A9, the results are also qualitatively similar if we compute \textit{Global R\text{-zone}} using a GDP-weighted average across countries. While still highly significant, the \( t \)-statistics on \textit{Global R\text{-zone}} are typically somewhat smaller when we use the GDP-weighted version.

---

8 As one would expect, relative to the equal-weighted definition of \textit{Global R\text{-zone}} shown in Table 7, defining \textit{Global R\text{-zone}} in this “leave-one-out” fashion tends to slightly raise the coefficient on \textit{Local R\text{-zone}} and slightly reduce the coefficient on \textit{Global R\text{-zone}}.
Table A1: Equity Indices

This table presents an overview of the equity indices used in our analysis. The data is retrieved from 4 sources: Global Financial Data (GFD), the International Monetary Funds (IMF) *International Financial Statistics*, Bloomberg and the Jordá, Schularick and Taylor MacroHistory database (JST).

<table>
<thead>
<tr>
<th>Country</th>
<th>Years</th>
<th>Source</th>
<th>Equity Index</th>
</tr>
</thead>
<tbody>
<tr>
<td>Argentina</td>
<td>1950-2018</td>
<td>GFD</td>
<td>Buenos Aires SE General Index (IVBNG)†</td>
</tr>
<tr>
<td>Australia</td>
<td>1950-2018</td>
<td>GFD</td>
<td>Australia ASX All-Ordinaries (w/GFD extension)</td>
</tr>
<tr>
<td>Austria</td>
<td>1950-2018</td>
<td>GFD</td>
<td>Austria Wiener Boersekreammer Share Index (WBKI)</td>
</tr>
<tr>
<td>Belgium</td>
<td>1950-2018</td>
<td>GFD</td>
<td>Brussels All-Share Price Index (w/GFD extension)</td>
</tr>
<tr>
<td>Brazil</td>
<td>1950-2018</td>
<td>GFD</td>
<td>GFD Indices Brazil Bolsa de Valores de Sao Paulo (Bovespa)†</td>
</tr>
<tr>
<td>Canada</td>
<td>1950-2018</td>
<td>GFD</td>
<td>Canada S&amp;P/TSX 300 Composite (w/GFD extension)</td>
</tr>
<tr>
<td>Chile</td>
<td>1975-2001</td>
<td>GFD</td>
<td>Santiago SE Indice de Precios Selectivos Acciones</td>
</tr>
<tr>
<td>Chile</td>
<td>1999-2018</td>
<td>IMF</td>
<td>Selective Price Index (IPSA)</td>
</tr>
<tr>
<td>Colombia</td>
<td>2001-2018</td>
<td>IMF</td>
<td>Index of prices on the Bogotá Stock Exchange</td>
</tr>
<tr>
<td>Czech Republic</td>
<td>1997-2018</td>
<td>IMF</td>
<td>PX-50 index</td>
</tr>
<tr>
<td>Denmark</td>
<td>1950-2018</td>
<td>GFD</td>
<td>OMX Copenhagen All-Share Price Index</td>
</tr>
<tr>
<td>Finland</td>
<td>1950-2018</td>
<td>GFD</td>
<td>OMX Helsinki All-Share Price Index</td>
</tr>
<tr>
<td>France</td>
<td>1950-1989</td>
<td>JST</td>
<td>Stock prices (nominal index)</td>
</tr>
<tr>
<td>France</td>
<td>1987-2018</td>
<td>GFD</td>
<td>Paris CAC-40 Index</td>
</tr>
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<td>1950-1961</td>
<td>JST</td>
<td>Stock prices (nominal index)</td>
</tr>
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<td>1952-2018</td>
<td>GFD</td>
<td>Athens SE General Index (w/GFD extension)</td>
</tr>
<tr>
<td>Hong Kong</td>
<td>1964-2018</td>
<td>GFD</td>
<td>Hong Kong Hang Seng Composite Index (w/GFD Extension)</td>
</tr>
<tr>
<td>Hungary</td>
<td>1994-2018</td>
<td>GFD</td>
<td>Vienna OETEB Hungary Traded Index (Forint)</td>
</tr>
<tr>
<td>Iceland</td>
<td>2002-2018</td>
<td>IMF</td>
<td>Index of the 15 largest and most traded Icelandic companies of the OMX</td>
</tr>
<tr>
<td>India</td>
<td>1950-2018</td>
<td>GFD</td>
<td>Bombay SE Sensitive Index (w/GFD extension)</td>
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<tr>
<td>Indonesia</td>
<td>1977-2018</td>
<td>GFD</td>
<td>Jakarta SE Composite Index</td>
</tr>
<tr>
<td>Ireland</td>
<td>1950-2018</td>
<td>GFD</td>
<td>Ireland ISEQ Overall Price Index (w/GFD extension)</td>
</tr>
<tr>
<td>Israel</td>
<td>1991-2019</td>
<td>Bloomberg</td>
<td>TA-125 (last price)</td>
</tr>
<tr>
<td>Italy</td>
<td>1950-2018</td>
<td>GFD</td>
<td>Banca Commerciale Italiana Index (w/GFD extension)</td>
</tr>
<tr>
<td>Japan</td>
<td>1950-1986</td>
<td>JST</td>
<td>Stock prices (nominal index)</td>
</tr>
<tr>
<td>Japan</td>
<td>1984-2017</td>
<td>GFD</td>
<td>Japan Nikkei 500 Index</td>
</tr>
<tr>
<td>Korea</td>
<td>1962-2018</td>
<td>GFD</td>
<td>Korea SE Stock Price Index (KOSPI)</td>
</tr>
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<td>Luxembourg</td>
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<td>Bloomberg</td>
<td>LUXXX Index (last price)</td>
</tr>
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<td>Mexico</td>
<td>1950-2018</td>
<td>GFD</td>
<td>Mexico SE Indice de Precio y Cotizaciones (IPC)</td>
</tr>
<tr>
<td>Netherlands</td>
<td>1950-2018</td>
<td>GFD</td>
<td>Netherlands All-Share Price Index (w/GFD extension)</td>
</tr>
<tr>
<td>New Zealand</td>
<td>1950-2018</td>
<td>GFD</td>
<td>New Zealand SE All-Share Capital Index</td>
</tr>
<tr>
<td>Norway</td>
<td>1950-1971</td>
<td>JST</td>
<td>Stock prices (nominal index)</td>
</tr>
<tr>
<td>Norway</td>
<td>1969-2018</td>
<td>GFD</td>
<td>Oslo SE All-Share Index</td>
</tr>
<tr>
<td>Peru</td>
<td>1988-2016</td>
<td>IMF</td>
<td>Share price index of the Lima Stock Exchange (industrials and mining)</td>
</tr>
<tr>
<td>Portugal</td>
<td>1950-2018</td>
<td>GFD</td>
<td>Oporto PSI-20 Index</td>
</tr>
<tr>
<td>Russia</td>
<td>1993-2018</td>
<td>GFD</td>
<td>Russia Moscow Index (MOEX) Composite</td>
</tr>
<tr>
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<td>1961-2018</td>
<td>GFD</td>
<td>Singapore FTSE Straits-Times Index</td>
</tr>
<tr>
<td>South Africa</td>
<td>1960-2018</td>
<td>IMF</td>
<td>All ordinary shares listed on Security Exchange South Africa</td>
</tr>
<tr>
<td>Spain</td>
<td>1950-1989</td>
<td>JST</td>
<td>Stock prices (nominal index)</td>
</tr>
<tr>
<td>Spain</td>
<td>1987-2018</td>
<td>GFD</td>
<td>Madrid SE IBEX-35</td>
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<td>GFD</td>
<td>Sweden OMX Affarsvarldens General Index</td>
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<td>Switzerland</td>
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<td>GFD</td>
<td>Switzerland Price Index (w/GFD extension)</td>
</tr>
<tr>
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<td>GFD</td>
<td>Thailand SET General Index</td>
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<td>Turkey</td>
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<td>GFD</td>
<td>Istanbul SE IMKB-100 Price Index</td>
</tr>
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<td>GFD</td>
<td>UK FTSE All-Share Index (w/GFD extension)</td>
</tr>
<tr>
<td>United States</td>
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<td>GFD</td>
<td>S&amp;P 500/Cowles Composite Price Index (w/GFD extension)</td>
</tr>
</tbody>
</table>

† Return index
Table A2: House Price Indices

This table presents an overview of the house price indices used in our analysis. The data is retrieved from 3 sources: Bank of International Settlements’ (BIS) Property Price Statistics, the OECD’s Household Prices database and the Jordá, Schularick and Taylor MacroHistory database (JST).

<table>
<thead>
<tr>
<th>Country</th>
<th>Years</th>
<th>Source</th>
<th>Variable</th>
</tr>
</thead>
<tbody>
<tr>
<td>Australia</td>
<td>1950-1972</td>
<td>JST</td>
<td>House prices (hpnom) normalized by consumer price index (cpi)</td>
</tr>
<tr>
<td>Australia</td>
<td>1970-2018</td>
<td>BIS</td>
<td>Real residential property prices</td>
</tr>
<tr>
<td>Austria</td>
<td>2000-2018</td>
<td>BIS</td>
<td>Real residential property prices</td>
</tr>
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<td>Belgium</td>
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<td>JST</td>
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</tr>
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</tr>
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<td>BIS</td>
<td>Real residential property prices</td>
</tr>
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<td>JST</td>
<td>House prices (hpnom) normalized by consumer price index (cpi)</td>
</tr>
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<td>BIS</td>
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<td>JST</td>
<td>House prices (hpnom) normalized by consumer price index (cpi)</td>
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<td>BIS</td>
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</tr>
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<td>JST</td>
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</tr>
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<td>BIS</td>
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<td>1988-2017</td>
<td>OECD</td>
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<td>Real residential property prices</td>
</tr>
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<td>Singapore</td>
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<td>Real residential property prices</td>
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<td>1971-2018</td>
<td>BIS</td>
<td>Real residential property prices</td>
</tr>
<tr>
<td>Sweden</td>
<td>1950-1972</td>
<td>JST</td>
<td>House prices (hpnom) normalized by consumer price index (cpi)</td>
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<tr>
<td>Sweden</td>
<td>1970-2018</td>
<td>BIS</td>
<td>Real residential property prices</td>
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<td>BIS</td>
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<td>BIS</td>
<td>Real residential property prices</td>
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Table A3: Debt Sample Overview

This table presents an overview of the sources for business debt (Panel A) and household debt (Panel B) used in our analysis. The data is retrieved from 3 sources: the International Monetary Funds (IMF) Global Debt Database, the Jordá, Schularick and Taylor MacroHistory database (JST) and the Bank of International Settlements’ (BIS) Total credit statistics.

Panel A: Business Debt Sources

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<th>Country</th>
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<th>Source</th>
<th>Variable</th>
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### Panel B: Household Debt Sources

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The table presents the results of three crisis prediction models: 

\[ y_{i,t \rightarrow t+h} = a_i + \beta \times \text{High Debt Growth}_{i,t} + \delta \times \text{High Price Growth}_{i,t} + \gamma \times \text{R-Zone}_{i,t} + \epsilon_{i,t} \]

where \( y_{i,t \rightarrow t+h} \in \{\text{Crisis}_{i,t+1 \rightarrow t+h}, \text{Crisis}_{i,t+1 \rightarrow t+h} - \text{Crisis}_{i,t \rightarrow t+h}, \text{CrisisStart}_{i,t \rightarrow t+h}\} \). \( \text{CrisisStart}_{i,t} \) is an indicator variable equal to 1 if a crisis begins in year \( t \) in country \( i \). \( \text{Crisis}_{i,t+1 \rightarrow t+h} = \max\{\text{CrisisStart}_{i,t+1 \rightarrow t+h}, \ldots, \text{CrisisStart}_{i,t+1 \rightarrow t+h}\} \) is an indicator variable, which takes the value of 1 if a crisis has occurred in country \( i \) between year \( t+1 \) and \( t+h \). High Debt Growth \( \equiv 1\{\Delta_3(\text{Debt/GDP})_{i,t} > 80^{\text{th}} \text{ percentile}\} \) is an indicator variable which takes the value of 1 if 3-year debt growth is the in the highest quintile, while High Price Growth \( \equiv 1\{\Delta_3(\text{Price})_{i,t} > 66^{\text{th}} \text{ percentile}\} \) is an indicator variable which takes the value of 1 if 3-year price growth is in its highest tertile. The R-Zone variable is the intersection of high price growth and high debt growth: \( \text{R-Zone} \equiv \text{High Debt Growth} \times \text{High Price Growth} \). We run the regression on both the business sector, using business debt and equity prices to define the indicators (Panel A), and the household sector, using household debt and house prices to define the indicators (Panel B). In Panel A1 and B1 t-statistics are based on Driscoll and Kraay (1998) standard errors with lags of 0, 3, 5 and 6 years for prediction horizons 1, 2, 3 and 4 years, respectively. In Panel A2, A3, B2 and B3 t-statistics are based on standard errors clustered by year. *, ** and *** denote significance at the 10%, 5% and 1% level, respectively, using Kiefer and Vogelsang (2005) corrected p-values. Reported coefficients and \( R^2 \)'s are in percent.

### Panel A: Business Sector

**Panel A1: Predicting \( \text{Crisis}_{i,t+1 \rightarrow t+h} \) (Baseline Regression)**

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<th>( \delta )</th>
<th>t-statistic</th>
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**Panel A2: Predicting \( \text{Crisis}_{i,t+1 \rightarrow t+h} - \text{Crisis}_{i,t \rightarrow t+h-1} \)**

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**Panel A3: Predicting \( \text{Crisis}_{i,t+h} \)**

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**Multivariate Regressions**

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Panel B: Household Sector

Panel B1: Predicting Crisis<sub>t+1 to t+h</sub> (Baseline Regression)

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Panel B2: Predicting Crisis<sub>t+1 to t+h</sub> - Crisis<sub>t+1 to t+h-1</sub>

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Panel B3: Predicting Crisis<sub>t+h</sub>

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Multivariate Regressions

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Table A6: Bootstrapped P-Values and Bias Estimates

This table presents p-values for coefficient estimates from our main crisis prediction model at the 3-year horizon:

\[ Crisis_{i, t+1 \ to \ t+3} = a_i + \beta \times \text{High Debt Growth}_{it} + \delta \times \text{High Price Growth}_{it} + \gamma \times \text{R-Zone}_{it} + \epsilon_{it} \]

We show the p-values calculated using standard asymptotics, fixed-b asymptotics as in Kiefer and Vogelsang (2005) and p-values calculated using the block bootstrap procedure described in section D. For the bootstrap we draw 100,000 samples with the block sizes being drawn from a geometric distribution with success probability 0.125. We show the p-values for coefficient estimates obtained from univariate regressions (corresponding to columns (3.1), (3.2) and (3.4) in Table 4), and p-values for coefficient estimates obtained from multiple regressions (corresponding to column (3.3) in Table 4). All t-statistics are calculated using Driscoll and Kraay (1998) standard errors with 5 lags. Panel A and B present the results for the business sector and household sector, respectively.

### Panel A: Business Sector

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<th>Multivariate Regressions</th>
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</tr>
<tr>
<td>( P(&gt;</td>
<td>t</td>
<td>) ) fixed-b</td>
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<tr>
<td>( P(&gt;</td>
<td>t</td>
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### Panel B: Household Sector

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Table A7: Sensitivity of Main Crisis Prediction to Cutoff and Sample Period

The table presents coefficient estimates and t-statistics of our main crisis prediction model:

\[ Crisis_{i,t+1} to t+3 = a_i + \beta \times \text{High Debt Growth}_{it} + \delta \times \text{High Price Growth}_{it} + \gamma \times \text{R-Zone}_{it} + \epsilon_{it} \]

where \( Crisis_{i,t+1} to t+3 \) is an indicator variable, which takes the value of 1 if a crisis has occurred in country \( i \) within 3 years of time \( t \). High Debt Growth \( _{it} \equiv 1\{\Delta_i(D/GDP)_{it} > C_D\} \) is an indicator variable which takes the value of 1 if 3-year debt growth in country \( i \) is higher than \( C_D \), while High Price Growth \( _{it} \equiv 1\{\Delta_i \log(Price_{it}) > C_P\} \) is an indicator variable which takes the value of 1 if 3-year price growth in country \( i \) is above \( C_P \). The R-Zone variable is the intersection of high price growth and high debt growth: \( \text{R-Zone}_{it} \equiv \text{High Debt Growth}_{it} \times \text{High Price Growth}_{it} \). The indicator variables are defined using a range of cutoffs for debt growth (\( C_D \) varies across columns) and price growth (\( C_P \) varies across rows), and the model is tested with both a univariate (left) and a multiple regression specification (right).

We test the model on both our full sample, a pre-2000 sample where we exclude data after 1999 (last prediction year is 1996), and post-2000 sample where we exclude business sector data prior to 1997 (last prediction year is 2012). We run the regressions on both the business sector, using business debt and equity prices to define the indicators (Panel A), and the household sector, using household debt and house prices to define the indicators (Panel B). t-statistics are based on Driscoll and Kraay (1998) standard errors with 5 lags. *, ** and *** denote significance at the 10%, 5% and 1% level, respectively, using Kiefer and Vogelsang (2005) corrected p-values.

### Panel A: Business Sector

#### Univariate Regressions

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#### Multivariate Regressions

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<tr>
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<tr>
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#### Pre-2000 Sample

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#### Multivariate Regressions

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#### Post-2000 Sample

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<td>0.9 1.2 1.6 2.5 8.7***</td>
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</table>

---

**Note:** The table above contains the coefficient estimates and t-statistics for the main crisis prediction model across different cutoffs and sample periods. Each row represents a different specification, with the columns indicating the cutoff points for debt growth and price growth. The model is tested with both univariate and multivariate regression specifications, and t-statistics are based on Driscoll and Kraay (1998) standard errors with 5 lags. Significance levels are indicated with *, **, and *** for 10%, 5%, and 1%, respectively, using Kiefer and Vogelsang (2005) corrected p-values.
## Panel B: Household Sector

### Univariate Regressions

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<tr>
<td>Pre-2000 Sample</td>
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<tr>
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### Multivariate Regressions

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<td>Post-2000 Sample</td>
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<tr>
<td>2 5 8 11 14</td>
<td></td>
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</table>
The table presents the results of the regression model:

$$\text{Crisis}_{i,t+1} \to t+h = a_h + \gamma_{Bus,h} \times \text{Local R-Zone}_{Bus}^{it} + \xi_{Bus,h} \times \text{Global R-Zone}_{Bus}^{it} + \gamma^{HH,h} \times \text{Local R-Zone}_{HH}^{it} + \xi^{HH,h} \times \text{Global R-Zone}_{HH}^{it} + \epsilon_{it}$$

where Local R-Zone$_{Bus}^{it}$ is an indicator variable capturing episodes of high business debt growth and equity price growth, while Local R-Zone$_{HH}^{it}$ is an indicator variable capturing episodes of high household debt growth and house price growth. Global R-Zone$_{Bus}$ measures the fraction of countries excluding country $i$ in the business R-Zone at time $t$, while Global R-Zone$_{HH}$ measures the fraction of countries excluding country $i$ in the household R-Zone at time $t$. t-statistics are reported in the brackets and based on Driscoll and Kraay (1998) with lags of 0, 3, 5 and 6 years for prediction horizons 1, 2, 3 and 4 years, respectively. *, ** and *** denote significance at the 10%, 5% and 1% level, respectively, using Kiefer and Vogelsang (2005) corrected p-values. Reported coefficients and $R^2$'s are in percent.

<table>
<thead>
<tr>
<th>Dependent Variable</th>
<th>Crisis within 1 year</th>
<th>Crisis within 2 years</th>
<th>Crisis within 3 years</th>
<th>Crisis within 4 years</th>
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<tr>
<td></td>
<td>(1.1) (1.2) (1.3)</td>
<td>(2.1) (2.2) (2.3)</td>
<td>(3.1) (3.2) (3.3)</td>
<td>(4.1) (4.2) (4.3)</td>
</tr>
<tr>
<td>Local R-Zone$<em>{Bus}$ ($\gamma</em>{Bus,h}$)</td>
<td>4.0 [1.0] 1.7 [0.5] 9.8* [1.9] 7.0 [1.4] 23.4** [2.9] 19.5** [2.3] 23.5** [2.6] 18.8* [2.1]</td>
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<tr>
<td>Global R-Zone$<em>{Bus}$ ($\xi</em>{Bus,h}$)</td>
<td>53.7* [1.8] 47.1 [1.4] 86.9*** [4.0] 54.0* [1.9] 110.4*** [4.7] 72.6 [1.8] 101.9*** [5.6] 33.8 [1.3]</td>
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<tr>
<td>Global R-Zone$_{HH}$ ($\xi^{HH,h}$)</td>
<td>24.6 [1.4] 5.4 [0.8] 53.4*** [2.7] 29.9* [1.9] 72.8*** [4.8] 37.7** [2.5] 92.8*** [7.2] 72.6*** [5.0]</td>
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<tr>
<td>$R^2$ (within)</td>
<td>6.1 4.8 7.3 9.3 10.3 12.5 14.3 14.4 19.0 10.7 16.1 18.1</td>
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<tr>
<td>Observations</td>
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</table>
Table A9: Crisis Prediction with Global R-Zones (GDP weighted)

The table presents the results of the regression model:

\[
Crisis_{i,t+1}^{t+h} = \alpha_i^h + \gamma_{Bus,h} \times \text{Local R-Zone}_{i,t}^{Bus} + \xi_{Bus,h} \times \text{GDP Weighted Global R-Zone}_{i,t}^{Bus} + \gamma_{HH,h} \times \text{Local R-Zone}_{i,t}^{HH} + \xi_{HH,h} \times \text{GDP Weighted Global R-Zone}_{i,t}^{HH} + \epsilon_i^h
\]

where Local R-Zone_{i,t}^{Bus} is an indicator variable capturing episodes of high business debt growth and equity price growth, while Local R-Zone_{i,t}^{HH} is an indicator variable capturing episodes of high household debt growth and house price growth. GDP Weighted Global R-Zone_{i,t}^{Bus} measures the fraction of countries in the business R-Zone weighted by their GDP at a given point in time, while GDP Weighted Global R-Zone_{i,t}^{HH} measures the fraction of countries weighted by their GDP in the household R-Zone at a given point in time. t-statistics are reported in the brackets and based on Driscoll and Kraay (1998) with lags of 0, 3, 5 and 6 years for prediction horizons 1, 2, 3 and 4 years, respectively. *, ** and *** denote significance at the 10%, 5% and 1% level, respectively, using Kiefer and Vogelsang (2005) corrected p-values. Reported coefficients and \(R^2\)'s are in percent.

<table>
<thead>
<tr>
<th></th>
<th>Crisis within 1 year</th>
<th>Crisis within 2 years</th>
<th>Crisis within 3 years</th>
<th>Crisis within 4 years</th>
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<td>(1.1)</td>
<td>(1.2)</td>
<td>(1.3)</td>
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<td>[1.4]</td>
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<td>Weighted Global R-Zone_{Bus}</td>
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\[ R^2 \ (within) \] 4.2 2.7 5.9 9.4 6.5 11.4 12.9 11.0 17.3 11.3 13.9 18.3

Observations 1,258 1,107 1,084 1,258 1,107 1,084 1,258 1,107 1,084 1,258 1,107 1,084