APPENDIX TO:

INVESTMENT AND THE CURRENT ACCOUNT IN THE SHORT RUN AND THE LONG RUN

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Appendix

This appendix contains more information about our data set and describes our estimation methods. Details about our data set appear in section A.1. Estimates of the contemporaneous correlation of the first differences of the current account, ΔCA_t, and investment ΔI_t for the rest of the G$-$7 economies (*i.e.*, France, Germany, Italy, Japan, the U.K., and the U.S.) are found in section A.2. Section A.3 discusses reduced-form VAR (RFVAR) estimation procedures and results for the entire G$-$7. Unit root tests and the results of these tests are furnished in section A.4. Section A.5 outlines the way in which we estimate the structural VARs (SVARs) of section 4 of the paper. The relationship between the short run and long run responses of the current account to investment and common world shocks is discussed for the remainder of the G$-$7 in section A.6. The appendix finishes with section A.7 that provides estimates of the relevant SVAR coefficients, forecast error variance decompositions (FEVDs), and graphical evidence for the remainder of the G$-$7.

A.1 Data

The paper uses quarterly data that spans the period 1973.1$-$1995.4. All of our estimates are based on the 1975.1 $-$ 1995.4 sample. We require data earlier than 1975.1 for lags when computing ADF regressions and estimating VARs. Our definition of investment equals the sum of gross capital formation and the change in stocks (*i.e.* inventories) with one exception. The U.S. investment series includes gross capital formation of the government. This definition of investment is the same one Glick and Rogoff (1995) use. The source of the quarterly gross capital formation series is Datastream. In Datastream, this data appears in billions of constant local currency units and is seasonally adjusted at annual rates. Quarterly data for the change in stocks for the G$-$7 and the U.S. gross capital formation of the government series are found in the IFS data bank. The IFS provides this data in current local currency units, seasonally adjusted, at annual rates. We convert this nominal data to real data using the GDP price deflators of the G$-$7. We obtain the GDP price deflators from Datastream. Subsequent to making an adjustment to annual rates where appropriate, this completes the task of creating the investment series.

Datastream reports quarterly current account data in millions of current U.S. dollars for the G$-$7 with the exception of the current account series for Canada, Japan, and the U.K. In Datastream, the current account series for Japan appears at the monthly frequency in millions of U.S. dollars. We temporally aggregate the monthly data to form a quarterly series. The current account series for Canada and the U.K. are provided in millions of current local currency units. All of the current account data in Datastream are seasonally adjusted with the exception of the series for France and Italy. We apply the X$-$11 filter to seasonally adjust these series. To convert the current account series for France, Germany, Italy, and Japan to millions of constant local currency units, we apply the appropriate U.S. dollar/local currency exchange rate. The quarterly exchange rate data we employ is from the IFS and has been generously provided by Mick Devereux. Next, we use GDP deflators to create quarterly current account series in millions of constant local currency units. The final step produces current account series for the G−7 in billions of constant local currency units at annual rates.

Figures A1 and A2 contain time plots of the level of investment, I_t, and the level of the current account, CA_t, for the G−7 economies in alphabetical order. The most striking aspect of this data is the observed inverse relationship between I_t and the CA_t. However, no structural interpretation can be given to this observation.

A.2 Reduced Form Univariate Regressions

At the beginning of section 3 of the paper, we present evidence that the contemporaneous correlation of ΔCA_t and ΔI_t for Canada matches estimates reported elsewhere. In this section, we present evidence that our entire G−7 data set produces similar estimates of the slope coefficient of the regression

$$\Delta CA_t \quad = \quad b_0 \quad + \quad b_1 \Delta I_t \quad + \quad \upsilon_t .$$

Among others, Glick and Rogoff (1995) note that much of the literature that studies the degree of international capital mobility examines the slope coefficient of this regression. Since Glick and Rogoff report estimates of b_1 that are less than zero, the estimates make dubious the Feldstein and Horioka (1980) story of autarkic national capital markets. However, at face value the Glick and Rogoff estimates cannot be taken as evidence of perfect capital mobility internationally because these estimates of b_1 are less negative than negative one.

The upper half of table $A1$ reproduces Glick and Rogoff's ordinary least squares (OLS) estimates. Our results appear in the bottom half of the table. The sample period for our regression begins with 1975.1 and ends at 1995.4, while the Glick-Rogoff estimates are based on annual data from 1975 to 1990.[A.1] For our estimates of b_1, we provide both OLS and Newey-West (1994) standard errors. Our estimates of b_1 are all negative with an average estimate of −0.33. The average Glick-Rogoff estimate is about −0.39. With the exception of Germany, all of our estimates of b_1 are smaller (in absolute value) than those Glick and Rogoff report. All of our estimates of b_1 possess $t-$ratios greater than two in absolute value.[A.2]

Thus, our estimates of the reduced form regressions indicate that the correlation between ΔCA_t and ΔI_t in our data set is similar to that found by other researchers. To reiterate our discussion just before section 3.1 of the paper, no structural interpretation can be given to these estimates without an identification scheme.

A.3 Reduced-Form VAR Estimation

The results of estimating the reduced form VAR of ΔCA_t and ΔI_t appear in tables $A2.1 - 2$. For each G−7 member, we compute OLS estimates of the fourth-order VAR

[A.1] Estimates for the 1975:1−1990:4 sample are qualitatively similar to those for the 1975:1−1995:4 sample.

[A.2] The major di®erence between our estimates and those of Glick and Rogo® is the R^2 and Durbin-Watson (D−W) statistics. For six of the seven regressions, the value of R^2 we report is smaller than those in the top panel of table A1. On the other hand, the Glick-Rogo® D−W statistics are smaller than ours. Most likely, the source of these di®erences is temporal aggregation.

$$(A3.1) \qquad \Delta I_t = A_{\Delta I,\,\Delta I}(\mathbf{L})\Delta I_{t-1} + A_{\Delta I,\,\Delta CA}(\mathbf{L})\Delta CA_{t-1} + \varepsilon_{\Delta I,t},$$

and

$$(A3.2) \qquad \Delta CA_t = A_{\Delta CA,\,\Delta I}(\mathbf{L})\Delta I_{t-1} + A_{\Delta CA,\,\Delta CA}(\mathbf{L})\Delta CA_{t-1} + \varepsilon_{\Delta CA,t}.$$

We estimate the reduced form VAR of (A3.1) and (A3.2) by ordinary least squares (OLS). These results appear in table $A2.1$. The Granger-causality tests we present in table $A2.1$ follow the advice of Hamilton (1994) and construct a test statistic that asymptotically possesses the χ^2 distribution with four degrees of freedom. To compute the forecast errors and shocks to the stochastic trends that appear in table $A2.2$, we estimate a slightly altered reduced form VAR. In this instance, we write equations (A3.1) and (A3.2) as

$$(A3.3) \qquad \Delta I_t = A_{\Delta I,\,\Delta I}(\mathbf{1})\Delta I_{t-1} + A_{\Delta I,\,\Delta I}(\mathbf{L})\Delta^2 I_{t-1}$$
$$+ A_{\Delta I,\,\Delta CA}(\mathbf{1})\Delta CA_{t-1} + A_{\Delta I,\,\Delta CA}(\mathbf{L})\Delta^2 CA_{t-1} + \varepsilon_{\Delta I,t},$$

and

$$(A3.4) \qquad \Delta CA_t = A_{\Delta CA,\,\Delta I}(\mathbf{1})\Delta I_{t-1} + A_{\Delta CA,\,\Delta I}(\mathbf{L})\Delta^2 I_{t-1}$$
$$+ A_{\Delta CA,\,\Delta CA}(\mathbf{1})\Delta CA_{t-1} + A_{\Delta CA,\,\Delta CA}(\mathbf{L})\Delta^2 CA_{t-1} + \varepsilon_{\Delta CA,t},$$

where the lag operators are of order $p-1$. Standard deviations of the forecast errors and their correlations are calculated from the OLS residuals of the regressions (A3.3) and (A3.4) for all of the G$-$7 economies. To generate the standard deviations of the stochastic trends and their correlation for the G$-$7, we combine the OLS estimates of the coefficients $A_{\Delta I,\,\Delta CA}(\mathbf{1})$, $A_{\Delta I,\,\Delta I}(\mathbf{1})$, $A_{\Delta CA,\,\Delta I}(\mathbf{1})$, and $A_{\Delta CA,\,\Delta CA}(\mathbf{1})$ with the covariance matrix of the reduced form OLS residuals of the regressions (A3.3) and (A3.4). We follow King and Watson (1997) and compute the standard errors of the stochastic trends and their correlation by the delta method.

In table $A2.1$, we report sums of the estimated coefficients of equations (A3.1) and (A3.2), their standard errors, LM tests, and Granger-causality tests. Although there are some noticeable patterns across the G$-$7 in the signs of these sums of coefficients, only those for Japan and the U.K. have a $t-$ratio greater than two. For Japan, the coefficient sums with a $t-$ratio greater than two are the autoregressive parameters of equations (A3.1) and (A3.2). For the U.K., the coefficient sums with $t-$ratios greater than two (in absolute terms) are in equation (A3.2), the ΔCA_t regression.

Results of the LM and Granger-causality tests produce a similar picture. The LM test computes the statistic $T \times R^2$ to provide information about the hypothesis that all of the slope coefficients of either equation (A3.1) or equation (A3.2) are jointly equal to zero. In this case, the test statistic is asymptotically distributed χ^2 with eight degrees of freedom. Of the 14 regressions, only four regressions, the ΔI_t regression for Japan and the ΔCA_t regressions for France, the U.K. and the U.S., reject the hypothesis at the five percent significance level. Likewise, the tests for Granger-causality suggest that ΔCA possesses no forecasting power for ΔI across the G$-$7. On the other hand, using equation (A3.2) to test the hypothesis that $A_{\Delta CA,\,\Delta I}(j) = 0,\ j = 1,\ \ldots,\ 4$, yields evidence that is a bit more mixed. In this case, ΔI possesses forecasting power for ΔCA for France, the U.K., and the U.S. at the five percent significance level or better. These Granger-causality tests lend some support for the notion that lags of ΔI_t do not matter for ΔCA_t, as implied by the intertemporal, small open economy model. There exists stronger evidence that lags of ΔCA_t do not predict movements in ΔI_t.[A.3] Taken together with the estimates of the coefficient sums, the results of the LM tests and the Granger-causality tests appear to support the inference that, except for Japan, ΔI_t is to a first approximation white noise, but that ΔCA_t is white noise for Canada, Germany, and Italy, at least, in the context of the RFVAR of (A3.1) and (A3.2).

The top panel of table 2.2 contains summary statistics of the one-step ahead forecast errors. Except for the U.K., the standard deviation of the innovation of the ΔI_t regression is greater than that for the ΔCA_t regression for all G$-$7 economies. In addition, all of the estimated standard deviations possess $t-$ratios greater than two. As expected, the contemporaneous correlation between the innovations of ΔI_t and ΔCA_t regressions is negative. Aside from the U.K., the absolute value of the $t-$ratios of these correlations is less than two. At short horizons, news about unrestricted forecasts of ΔI_t and ΔCA_t are orthogonal.

Estimates of the stochastic trends appear in the bottom panel of table 2.2. For the entire G$-$7, the standard deviation of the permanent component of the innovation of the ΔI_t regression is greater than the same statistic for the ΔCA_t regression. Among this set of standard deviations, the standard deviation of the permanent component of the innovation of the ΔI_t regression for the U.K. possesses the smallest $t-$ratio of 1.94. For each of the G$-$7, the contemporaneous correlation between the permanent innovations in the ΔI_t and ΔCA_t regressions is negative. However, only the $t-$ratios for Italy, the U.K., and the U.S. are greater than two (in absolute terms). This suggests that unrestricted long-run movements in I_t and CA_t have a common source.

A.4 Unit Root Tests

Elliot, Rothenberg, and Stock (1996) present a method to test for a unit root in the presence of a deterministic mean or trend that is asymptotically more powerful than the usual DF $t-$ratio. This method begins by estimating the regression $y_t = \beta_0 + \beta_1 t + \omega_t$. The next step constructs the predicted values of $\hat{\omega}_t = y_t - \hat{\beta}_0 - \hat{\beta}_1 t$ to use in the augmented DF regression

[A.3] She®rin and Woo (1990), Otto (1992) and Ghosh (1995) present similar Granger-causality results for these countries. However, the regressions these authors estimate use the CA_t instead of ¢CA_t.

$$\hat{\omega}_t \quad = \quad \rho\hat{\omega}_{t-1} \quad + \quad \sum_{j=1}^{k} \vartheta_j \Delta\hat{\omega}_{t-j} \quad + \quad u_{t,k},$$

where the lag length, k, is chosen using the BIC criterion to render u_t white noise.[A.4] The GLS-DF $t-$ratio is constructed using the OLS estimates of ρ and its standard error. Elliot, Rothenberg, and Stock provide asymptotic ten percent, five percent, and one percent critical values for the GLS-DF $t-$ratio equal to -2.57, -2.89, and -3.48, respectively.

We present OLS estimates of the standard ADF $t-$ratio using the regression

$$y_t \quad = \quad \psi_0 \; + \; \psi_1 t \; + \; \gamma y_{t-1} \; + \; \sum_{j=1}^{k} \xi_j \Delta y_{t-j} \; + \; e_t,$$

where the lag length, k, is chosen to render e_t white noise. The lag length, k, is chosen using the Campbell and Perron (1991) rule. This rule selects a maximum k *a priori* and then discards lags until the $t-$ratio of Δy_{t-k} becomes less than 1.6 (in absolute value). MacKinnon (1991) provides asymptotic ten percent, five percent, and one percent critical values for the DF $t-$ratio equal to -3.13, -3.41, and -3.96, respectively.

Unit root tests are often criticized because of power problems these tests have, for example, with trend stationary alternatives.[A.5] Another problem that face unit root tests is an inability to provide information about the sampling variability of the estimate of γ. Stock (1991) presents the results of Monte Carlo simulations that are the building blocks for the construction of asymptotic confidence intervals. We present 95 percent asymptotic confidence intervals using the ADF regression with an intercept and a linear trend. As Stock suggests, we use linear interpolation to construct the asymptotic confidence intervals of γ.

A.5 SVAR Estimation Methods

Our estimation strategy follows closely that of King and Watson (1997). Since the appendix King and Watson (1997) supply contains a large amount of detail about their estimation methods, this section of our appendix provides only a brief sketch of the way in which we adapt these estimation methods. In particular, this section includes a description of the methods we use to estimate the SVAR when either $\mathcal{LR}_{I,CA}$ or $\mathcal{LR}_{CA,I}$ serves as the identifying restriction.

To estimate a SVAR when a long run multiplier acts as the identifying restriction, King and Watson (1997) use the technique of rewriting a regression to include second difference terms. When $\mathcal{LR}_{I,CA}$ identifies the SVAR, we develop the bivariate system to estimate by writing equation (8) as

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[A.4] To choose the lag length, the BIC criterion, ln[$\hat{\sigma}^2_{u;k}$] + (k=T)ln[T], is minimized over k.

[A.5] These problems are at the center of the debate of the source of the trend in real U.S. GNP. Diebold and Senhadji (1996) and Nelson and Murray (1997) present contrasting views of this issue.

$$\text{(A5.1)} \qquad \Delta I_t \;=\; \lambda_{I,CA}(\mathbf{1})\Delta CA_t \;+\; \lambda_{I,I}(\mathbf{1})\Delta I_{t-1}$$

$$+\;\Lambda_{I,I}(\mathbf{L})\Delta^2 I_{t-1} \;+\; \Lambda_{I,CA}(\mathbf{L})\Delta^2 CA_{t-1} \;+\; \eta_{W,t},$$

where, for example,

$$\Lambda_{I,I}(\mathbf{L})\Delta^2 I_{t-1} \;=\; -\sum_{j=1}^{p-1}\left(\sum_{s=j+1}^{p}\lambda_{I,I,s}\right)\Delta^2 I_{t-j}.$$

From (8), it follows that $\mathcal{LR}_{I,CA} = \lambda_{I,CA}(\mathbf{1})/[1 - \lambda_{I,I}(\mathbf{1})]$. With a bit of algebra, this yields the regression

$$\text{(A5.2)} \quad \Delta I_t \;-\; \mathcal{LR}_{I,CA}\Delta CA_t \;=\; \lambda_{I,I}(\mathbf{1})[\Delta I_{t-1} \;-\; \mathcal{LR}_{I,CA}\Delta CA_t]$$

$$+\;\Lambda_{I,I}(\mathbf{L})\Delta^2 I_{t-1} \;+\; \Lambda_{I,CA}(\mathbf{L})\Delta^2 CA_{t-1} \;+\; \eta_{W,t}.$$

Since ΔCA_t can be correlated with $\eta_{W,t}$, we compute the coefficients of this equation with an IV estimator using the instruments $\{\Delta I_{t-j},\ \Delta CA_{t-j}\}_{j=1}^{p}$ for the G$-$7 economies. With the coefficient estimates of the regression of (A5.2) in hand, we estimate equation (9) by IV with the instruments $\{\Delta I_{t-j},\ \Delta CA_{t-j}\}_{j=1}^{p}$ and $\hat{\eta}_{W,t}$. The instrument $\hat{\eta}_{W,t}$ denotes the residuals of regression (A5.2). We use a symmetric procedure when $\mathcal{LR}_{CA,I}$ serves to identify the SVAR. Across the G$-$7, the standard errors of the estimates of $\mathcal{LR}_{I,CA}$ and $\mathcal{LR}_{CA,I}$ are computed using the delta method.

Another issue we face is that the procedure just described uses a generated regressor, $\hat{\eta}_{W,t}$, as an instrument to estimate equation (9). The generated regressor problem arises because the variables on the right hand side of equation (9) and $\hat{\eta}_{W,t}$ may be correlated. As a result, an adjustment is needed to the estimator of the covariance matrix of the coefficients of equation (9). King and Watson (1997) present the details of the adjustment to this covariance matrix. We do not duplicate their efforts here.

To compute the empirical standard errors of the FEVDs we report in tables 3 and $A5.1-3$, the Monte Carlo procedure begins with estimates of the intercepts, slope coefficients, and the covariance matrix of the residuals of the RFVAR of equations (A3.1) and (A3.2) for each member of the G$-$7. Using these reduced-form estimates, we generate 1000 pairs of normally distributed, mean zero random variates. The covariance matrix of these random variates equals the covariance matrix of the reduced-form residuals. From these synthetic residuals and the estimated intercepts and slope coefficients we build up artificial I_t and the CA_t series. Next, we estimate the SVARs under the identifications of $R1$, $R2$, and $R6$ using these 1000 replications. The estimates of the SVAR using the artificial data allows us to construct the small sample standard errors of the FEVDs. Since this procedure builds the Monte Carlo up from the distribution of the RFVAR residuals, the standard errors of the FEVDs possesses only an interpretation as draws from the small sample or empirical distribution of the joint dynamic process that generates I_t and the CA_t.

A.6 How Do $\lambda_{\mathsf{I;CA;0}}$, $\lambda_{\mathsf{CA;I;0}}$, and $\mathcal{LR}_{\mathsf{I;CA}}$ Affect Estimates of $\mathcal{LR}_{\mathsf{CA;I}}$?

To show that $\mathcal{LR}_{CA,I}$ depends on the value of $\lambda_{CA,I,0}$, note that the reduced-form VAR (A3.1) and (A3.2)

$$\begin{bmatrix} I_t \\ CA_t \end{bmatrix} = \begin{bmatrix} A_{\Delta I,\,\Delta I}(\mathbf{L}) & A_{\Delta I,\,\Delta CA}(\mathbf{L}) \\ A_{\Delta CA,\,\Delta I}(\mathbf{L}) & A_{\Delta CA,\,\Delta CA}(\mathbf{L}) \end{bmatrix} \begin{bmatrix} I_{t-1} \\ CA_{t-1} \end{bmatrix} + \begin{bmatrix} \varepsilon_{\Delta I,t} \\ \varepsilon_{\Delta CA,t} \end{bmatrix},$$

implies that

$$\begin{bmatrix} \varepsilon_{\Delta I,t} \\ \varepsilon_{\Delta CA,t} \end{bmatrix} = \begin{bmatrix} 1 & -\lambda_{I,CA,0} \\ -\lambda_{CA,I,0} & 1 \end{bmatrix}^{-1} \begin{bmatrix} \eta_{W,t} \\ \eta_{C,t} \end{bmatrix},$$

or equation-by-equation $\varepsilon_{\Delta I,t} = [1 - \lambda_{I,CA,0}\lambda_{CA,I,0}]^{-1}[\eta_{W,t} + \lambda_{I,CA,0}\eta_{C,t}]$ and $\varepsilon_{\Delta CA,t} = [1 - \lambda_{I,CA,0}\lambda_{CA,I,0}]^{-1}[\lambda_{CA,I,0}\eta_{W,t} + \eta_{C,t}]$. The former expression reveals that movements in $\eta_{C,t}$ become less important for fluctuations in ΔI_t as $\lambda_{I,CA,0}$ goes toward zero. In symmetric fashion, the expression for $\varepsilon_{\Delta CA,t}$ shows that fluctuations in $\eta_{W,t}$ matter more for ΔCA_t as $\lambda_{CA,I,0}$ moves from zero toward negative one. Next, construct the long-run trends of I_t and CA_t

$$\mathcal{LR}_{I,t} = [1 - A_{\Delta I,\,\Delta I}(\mathbf{1})]^{-1}[A_{\Delta I,\,\Delta CA}(\mathbf{1})\mathcal{LR}_{CA,t} + \varepsilon_{\Delta I,t}],$$

and

$$\mathcal{LR}_{CA,t} = [1 - A_{\Delta CA,\,\Delta CA}(\mathbf{1})]^{-1}[A_{\Delta CA,\,\Delta I}(\mathbf{1})\mathcal{LR}_{I,t} + \varepsilon_{\Delta CA,t}].$$

Substituting $\varepsilon_{\Delta I,t}$ and $\varepsilon_{\Delta CA,t}$ from above into these expressions, and doing a bit of algebra, it is straightforward to show

$$\text{(A5.3)} \qquad \frac{\partial\mathcal{LR}_{I,t+j}/\partial\eta_{C,t}}{\partial\mathcal{LR}_{CA,t+j}/\partial\eta_{C,t}} = \frac{\lambda_{I,CA,0}[1 - A_{\Delta CA,\,\Delta CA}(\mathbf{1})] + A_{\Delta I,\,\Delta CA}(\mathbf{1})}{\lambda_{I,CA,0}A_{\Delta CA,\,\Delta I}(\mathbf{1}) + [1 - A_{\Delta I,\,\Delta I}(\mathbf{1})]},$$

and

$$\text{(A5.4)} \qquad \frac{\partial\mathcal{LR}_{CA,t+j}/\partial\eta_{W,t}}{\partial\mathcal{LR}_{I,t+j}/\partial\eta_{W,t}} = \frac{\lambda_{CA,I,0}[1 - A_{\Delta I,\,\Delta I}(\mathbf{1})] + A_{\Delta CA,\,\Delta I}(\mathbf{1})}{\lambda_{CA,I,0}A_{\Delta I,\,\Delta CA}(\mathbf{1}) + [1 - A_{\Delta CA,\,\Delta CA}(\mathbf{1})]}.$$

The long-run derivatives (A5.3) and (A5.4) are equivalent to $\mathcal{LR}_{I,CA}$ and $\mathcal{LR}_{CA,I}$, respectively. Since $\lim_{j\to\infty}\partial X_{t+j}/\partial\eta_{W,t} = \partial\mathcal{LR}_{X,t}/\partial\eta_{W,t}$ and $\lim_{j\to\infty}\partial X_{t+j}/\partial\eta_{C,t} = \partial\mathcal{LR}_{X,t}/\partial\eta_{C,t}$, where $X_t = I_t,\ CA_t$, we can equate the left hand sides of (A5.3) and (A5.4) with the response of I_t to a permanent movement in CA_t and to the response of CA_t to a permanent movement in I_t, respectively.

We use the derivatives (A5.3) and (A5.4) to study the effect of the different identification schemes on estimates of $\mathcal{LR}_{CA,I}$. First, consider imposing $R3$ on (A5.3) which yields $\mathcal{LR}_{I,CA}(R3) = [1 - A_{\Delta I,\,\Delta I}(\mathbf{1})]^{-1}A_{\Delta I,\,\Delta CA}(\mathbf{1})$. Second, evaluate (A5.4) at $R4$ to produce $\mathcal{LR}_{CA,I}(R4) = [1 - A_{\Delta CA,\,\Delta CA}(\mathbf{1})]^{-1}A_{\Delta CA,\,\Delta I}(\mathbf{1})$. These long-run multipliers, $\mathcal{LR}_{I,CA,I}(R3)$ and $\mathcal{LR}_{CA,I}(R4)$, together with a bit of algebra, allows us to write the derivative of (A5.4) as

$$\text{(A5.5)}\mathcal{LR}_{CA,I} = \frac{\lambda_{CA,I,0} + \mathcal{LR}_{CA,I}(R4)[1 - A_{\Delta I,\,\Delta I}(\mathbf{1})]^{-1}[1 - A_{\Delta CA,\,\Delta CA}(\mathbf{1})]}{\lambda_{CA,I,0}\mathcal{LR}_{I,CA}(R3) + [1 - A_{\Delta I,\,\Delta I}(\mathbf{1})]^{-1}[1 - A_{\Delta CA,\,\Delta CA}(\mathbf{1})]},$$

where we use $\mathcal{LR}_{CA,I} \equiv [\partial\mathcal{LR}_{CA,t+j}/\partial\eta_{W,t}]/[\partial\mathcal{LR}_{I,t+j}/\partial\eta_{W,t}]$.

Equation (A5.5) shows how our assumptions about $\lambda_{CA,I,0}$ drive point estimates of $\mathcal{LR}_{CA,I}$. The second term in the numerator, $\mathcal{LR}_{CA,I}(R4)[1-A_{\Delta I,\Delta I}(\mathbf{1})]^{-1}[1-A_{\Delta CA,\Delta CA}(\mathbf{1})]$, equals to -0.03, 0.18, -0.17, -0.02, -0.17, and -0.14 for France, Germany, Italy, Japan, the U.K., and the U.S., respectively. Thus, when $\lambda_{CA,I,0}$ (the only other term in the numerator) is close to zero, the numerator itself is close to zero. As $\lambda_{CA,I,0}$ becomes smaller than, say -0.35, the numerator of (A5.5) takes on the sign (negative), and approximately the value, of $\lambda_{CA,I,0}$. In the denominator, the term $[1 - A_{\Delta I,\,\Delta I}(\mathbf{1})]^{-1}[1 - A_{\Delta CA,\,\Delta CA}(\mathbf{1})]$ dominates. Since $[1 - A_{\Delta I,\,\Delta I}(\mathbf{1})]^{-1}[1 - A_{\Delta CA,\,\Delta CA}(\mathbf{1})]$ takes on the value of 0.46, 1.21, 0.79, 0.86, 0.39, and 1.05 for France, Germany, Italy, Japan, the U.K., and the U.S., respectively, this term is greater than $\lambda_{CA,I,0}\mathcal{LR}_{I,CA}(R3)$ for any value we choose to impose on $\lambda_{CA,I,0}$ to identify the SVAR.

A.7 SVAR Results for the G-7 sans Canada

Tables $A4.1-3$ contain point estimates of $\lambda_{I,CA,0}$, $\lambda_{CA,I,0}$, $\mathcal{LR}_{I,CA}$, and $\mathcal{LR}_{CA,I}$ as well as Wald statistics that test a key prediction of the intertemporal model under restrictions $R1-R6$ for the G$-$7 minus Canada. The point estimates we present in these tables reinforce the results for Canada of table 2. These results are that

(a) estimates of $\mathcal{LR}_{I,CA}$ back the inference that only the common world shock, $\eta_{W,t}$, drive permanent fluctuations in I_t,

(b) estimates of the impact coefficients $\lambda_{I,CA,0}$ and $\lambda_{CA,I,0}$ are sensitive to the identification scheme,

(c) estimates of $\mathcal{LR}_{CA,I}$ are closely tied to the value of $\lambda_{CA,I,0}$, and

(d) rejections of the hypothesis of the intertemporal, small open economy prediction that movements in the CA_t are independent of $\eta_{W,t}$ and its lags depend on the choice of the identification scheme.

Support for item (a) arise in tables $A4.1-3$ because of the $t-$ratios the estimates of $\mathcal{LR}_{I,CA}$ imply. Of the 30 estimates of $\mathcal{LR}_{I,CA}$ that appear in these tables, only six yield a

$t-$ratio greater than two in absolute value (see the top panel of table $A4.1$ and the bottom panel of table $A4.2$). Italy and the U.K. produce two-thirds of these estimates.

Item (b) stands out clearly from an inspection of tables $A4.1-3$. For example, under $R1$ (the top panel of table $A4.1$) all the estimates of $\lambda_{I,CA,0}$ are negative and four of the six estimates possess $t-$ratios greater than two in absolute value. However, under $R2$ (the bottom panel of table $A4.1$) four of the six estimates of $\lambda_{I,CA,0}$ are positive and none of these estimates are statistically different from zero at any reasonable significance level.

We show in section 3.5 of the paper that estimates of $\mathcal{LR}_{CA,I}$ are tied either to the identification of $\lambda_{CA,I,0}$ or its point estimate. The estimates of $\mathcal{LR}_{CA,I}$ we report in tables $A4.1-3$ bolster this analysis. This observation marks the basis for item (c). Except for the identification scheme $R2$ (see the bottom panel of table $A4.1$), country-by-country estimates of $\mathcal{LR}_{CA,I}$ are quite close to the value of $\lambda_{CA,I,0}$ given an identification.

Casual inspection of the Wald statistics of tables $A4.1-3$ is enough to clinch support for item (d). The Wald statistics that identifications $R1$ and $R3$ generate make this plain. Under $R1$ the null hypothesis of (11) is only rejected by the French data (see the top panel of table $A4.1$). However, this hypothesis receives strong rejections by data from France, Germany, Italy, the U.K., and the U.S. under $R3$ (see the top panel of table $A4.2$). The Wald statistics continue to back our thesis that tests of many of the predictions of the intertemporal model yield inferences sensitive to the identification.

The forecast error variance decompositions (FEVD) lend support to item (a). When $\mathcal{LR}_{I,CA} = 0$ is not imposed as the identifying restriction (*i.e.*, $R2$), we find that the common world shock $\eta_{W,t}$ explains more than 65 percent of the variation in I_t at a 24 quarter forecast horizon in eight of the remaining 12 FEVDs. In the case of $R2$, the FEVDs of I_t reveals that anywhere from 95 to 100 percent can be attributed to $\eta_{W,t}$ at impact (see the top panel of table $A5.2$). The FEVDs of I_t for France, Germany, Italy, Japan, the U.K., and the U.S. that we report in the top panels of tables $A5.1-3$ provide economically meaningful evidence that a shock common to the G$-$7 contributes most to fluctuation in I_t. This evidence is particularly strong at longer forecast horizons.

Tables $A5.1-3$ also contain results that help to sustain our claim that the persistence in the CA_t of the G$-$7 is an important and neglected aspect of the intertemporal approach to the current account. Under $R1$ (see the bottom panel of table $A5.1$), country-specific shocks in the form of $\eta_{C,t}$ generate about three-fourths or more of movements in the CA_t from impact to a forecast horizon of one year for the G$-$7 minus Canada. Since $R1$ restricts long-run fluctuations in the CA_t to respond only to $\eta_{C,t}$, it is expected that the analogous FEVDs at the longer forecast horizons should approach 100 percent. In this regard, the FEVDs in the bottom half of table $A5.2$ are particularly striking. These FEVDs are calculated under a long-run identifying restriction imposed on I_t, $R2$. Given no restriction on the behavior of the CA_t, we find that about 60 percent or more of the variation in this variable is explained by $\eta_{C,t}$ at all forecast horizons for all six economies. The reduced-form long-run identification of $R6$ reverses this result. The bottom half of table $A5.3$ contains only one FEVD that possesses a $t-$ratio greater than two (see the impact FEVD for the U.K.) and none of these

FEVDs is greater than 32 percent.

Our analysis of section 3.5 of the paper resolves the disparate results of the FEVDs of the CA_t we report in the bottom panels of tables $A5.1 - 3$. This analysis outlines the close connection between identifications that impose a restriction on the behavior of CA_t within the SVAR of (8) and (9) and the estimated unrestricted behavior of the CA_t. As a result, we should expect to observe FEVDs in which only $\eta_{C,t}$ should matter for the CA_t under $R1$. The FEVDs of the bottom panel of table $A5.1$ back this expectation. Likewise, under $R6$, we should anticipate that only common world shocks are responsible for fluctuations in the CA_t as we observe in the bottom panel of table $A5.3$. Hence, the FEVDs of the CA_t under $R2$ (see the bottom panel of table $A5.2$) are a strong signal for the importance of $\eta_{C,t}$ and the impact of these shocks on the persistence of the CA_t.

Figures $A3 - 8$ and $A9 - 14$ replicate figures 1 and 2 of the paper, respectively. By and large, the information that figures $A3-8$ and $A9-14$ contain reinforce the discussion we present in the paper for figures 1 and 2. That is, the 95 percent confidence intervals of figures $A3 - 8$ show that the identification matters for inference about $R1$ and $R6$. These figures are in line with figure 1 because support for the restriction of $R6$ exists only when $\lambda_{CA,I,0}$ is close to negative one. The 95 percent confidence ellipses of figures $A9 - 14$ provide more evidence to back our conclusion that minor changes to the identification leads to different views of the efficacy of the intertemporal, small open economy model.

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Table A1. Reduced Form Univariate Regressions

$$\Delta CA_t = b_0 + b_1 \Delta I_t + \upsilon_t$$

Glick and Rogoff Results
Sample Period: 1975 – 1990

	Canada	France	Germany	Italy	Japan	U.K.	U.S.
$\hat{b}_1$	-0.30	-0.42	-0.23	-0.59	-0.37	-0.60	-0.20
	(0.10)	(0.12)	(0.19)	(0.10)	(0.10)	(0.10)	(0.09)
R^2	0.38	0.47	0.10	0.71	0.49	0.72	0.27
D−W	1.88	1.42	1.60	1.19	1.54	1.81	1.18

Results Using Quarterly Data
Sample Period: 1975.1 – 1995.4

	Canada	France	Germany	Italy	Japan	U.K.	U.S.
$\hat{b}_1$	-0.37	-0.37	-0.33	-0.43	-0.20	-0.54	-0.11
	(0.08)	(0.09)	(0.09)	(0.09)	(0.08)	(0.11)	(0.04)
	[0.07]	[0.09]	[0.15]	[0.11]	[0.06]	[0.09]	[0.04]
R^2	0.20	0.17	0.13	0.22	0.07	0.22	0.09
D−W	2.34	2.46	2.50	2.36	1.64	2.72	2.20

The Glick and Rogoff (1995) estimates, taken from the bottom panel of their table 1, are based on annual data. We use OLS to compute estimates of the slope coefficient b_1 as do Glick and Rogoff. OLS standard errors appear in parenthesis. The values in brackets are Newey-West corrected standard errors. The Newey-West standard errors are constructed using an automatic lag length adjustment.

Table A2.1 Reduced Form VARs

Sample Period: 1975.1 – 1995.4

Coefficient Sums and Tests of Predictive Content

	Canada	France	Germany	Italy	Japan	U.K.	U.S.
$A_{\Delta I,\,\Delta I}(\mathbf{1})$	0.07	0.31	-0.16	0.00	0.46	0.18	-0.09
	(0.23)	(0.20)	(0.28)	(0.22)	(0.14)	(0.30)	(0.24)
$A_{\Delta I,\,\Delta CA}(\mathbf{1})$	0.16	0.30	-0.28	0.05	-0.03	-0.02	-0.26
	(0.37)	(0.34)	(0.33)	(0.30)	(0.22)	(0.36)	(0.72)
$A_{\Delta CA,\,\Delta I}(\mathbf{1})$	-0.02	-0.10	0.14	-0.26	-0.02	-0.93	-0.13
	(0.19)	(0.17)	(0.26)	(0.22)	(0.12)	(0.32)	(0.09)
$A_{\Delta CA,\,\Delta CA}(\mathbf{1})$	-0.42	-0.51	0.04	-0.26	0.37	-1.12	-0.04
	(0.30)	(0.29)	(0.31)	(0.29)	(0.18)	(0.38)	(0.26)
ΔI Regression : R^2	0.09	0.11	0.15	0.15	0.36	0.03	0.09
	[0.44]	[0.34]	[0.12]	[0.14]	[0.00]	[0.95]	[0.44]
ΔCA Regression : R^2	0.11	0.19	0.07	0.06	0.10	0.20	0.21
	[0.31]	[0.04]	[0.66]	[0.76]	[0.42]	[0.03]	[0.02]
Wald Test: $\Delta CA \rightarrow \Delta I$:	2.77	3.46	7.40	6.73	5.81	0.49	0.36
	[0.60]	[0.48]	[0.12]	[0.15]	[0.21]	[0.97]	[0.99]
Wald Test: $\Delta I \rightarrow \Delta CA$:	6.75	11.05	1.29	3.72	1.37	10.46	9.51
	[0.15]	[0.03]	[0.86]	[0.45]	[0.85]	[0.03]	[0.05]

Standard errors appear in parenthesis. The values in brackets are p-values. The p-values that appear below the R^2s represent significance levels for the LM test statistic $T \times R^2$. These statistics are asymptotically distributed χ^2 with eight degrees of freedom. The Wald statistics that tests the predictive power of either ΔI for ΔCA or ΔCA for ΔI possess a χ^2 distribution with four degrees of freedom asymptotically.

Table A2.2 Reduced Form VARs

Sample Period: 1975.1 – 1995.4

Forecast Errors

	Canada	France	Germany	Italy	Japan	U.K.	U.S.
$\sigma(\varepsilon_{\Delta I})$	4.40	21.64	22.25	10716.18	1639.38	4.22	58.04
	(1.03)	(5.06)	(5.20)	(2502.97)	(382.91)	(0.98)	(13.56)
$\sigma(\varepsilon_{\Delta CA})$	3.60	18.40	20.75	10332.76	1312.30	4.45	21.05
	(0.84)	(4.30)	(4.85)	(2413.41)	(306.51)	(1.04)	(4.92)
$Corr(\varepsilon_{\Delta I},\ \varepsilon_{\Delta CA})$	-0.44	-0.40	-0.46	-0.49	-0.23	-0.51	-0.35
	(0.27)	(0.28)	(0.26)	(0.25)	(0.31)	(0.24)	(0.29)

Shocks to Stochastic Trends

	Canada	France	Germany	Italy	Japan	U.K.	U.S.
$\sigma(\varepsilon_{I})$	4.54	28.79	21.32	10409.73	3096.24	5.26	56.84
	(1.49)	(10.80)	(5.28)	(3741.63)	(1279.11)	(2.71)	(17.22)
$\sigma(\varepsilon_{CA})$	2.54	12.86	19.79	9369.82	2104.67	3.85	24.11
	(0.67)	(3.01)	(8.67)	(2269.92)	(889.18)	(0.51)	(10.81)
$Corr(\varepsilon_{I},\ \varepsilon_{CA})$	-0.37	-0.39	-0.54	-0.63	-0.30	-0.88	-0.65
	(0.38)	(0.45)	(0.35)	(0.26)	(0.45)	(0.28)	(0.26)

Standard errors appear in parenthesis.

Table A3. Unit Root Tests

Sample Period: 1975.1 – 1995.4

Investment

	Canada	France	Germany	Italy	Japan	U.K.	U.S.
DF-GLS $t-$ratio	−2.64	−1.80	−1.48	−3.88	−2.56	−1.82	−3.37
Lag Length	1	1	1	2	3	1	1
ADF $t-$ratio	−2.78	−2.40	−1.67	−4.07	−1.99	−2.11	−3.90
Lag Length	1	2	0	2	3	0	3
ADF AR Root	0.85	0.90	0.93	0.74	0.96	0.90	0.73
95% CI: LL	0.73	0.79	0.89	−−	0.85	0.83	−−
95% CI: UL	1.04	1.05	1.06	0.90	1.05	1.05	0.93

Current Account

	Canada	France	Germany	Italy	Japan	U.K.	U.S.
DF-GLS $t-$ratio	−1.72	−2.81	−1.41	−1.80	−2.41	−2.15	−1.88
Lag Length	1	1	1	1	2	1	1
ADF $t-$ratio	−2.00	−3.41	−1.86	−2.30	−2.39	−2.57	−1.84
Lag Length	0	0	4	0	2	1	3
ADF AR Root	0.87	0.75	0.91	0.86	0.91	0.85	0.92
95% CI: LL	0.85	0.61	0.86	0.80	0.79	0.76	0.87
95% CI: UL	1.05	1.02	1.05	1.05	1.05	1.04	1.05

The asymptotic ten percent, five percent, and one percent critical values for the GLS-DF (ADF) $t-$ratio equal $-2.57\,(-3.13)$, $-2.89\,(-3.41)$, and $-3.48\,(-3.96)$, respectively.

Table A4.1 Structural VARs

Sample Period: 1975.1 – 1995.4

$R1: \mathcal{LR}_{\mathsf{CA;I}} = 0$

	France	Germany	Italy	Japan	U.K.	U.S.
$\lambda_{I,CA,0}$	-0.64	-0.37	-0.70	-0.33	-0.97	-1.68
	(0.30)	(0.24)	(0.19)	(0.34)	(0.23)	(0.58)
$\lambda_{CA,I,0}$	0.15	-0.12	0.26	0.03	1.15	0.12
	(0.28)	(0.21)	(0.25)	(0.22)	(0.71)	(0.09)
$\mathcal{LR}_{I,CA}$	-0.88	-0.58	-0.70	-0.44	-1.21	-1.54
	(0.66)	(0.32)	(0.25)	(0.54)	(0.20)	(0.59)
Wald Statistic	11.64	1.48	2.86	6.80	3.69	7.24
	[0.04]	[0.92]	[0.72]	[0.24]	[0.59]	[0.20]

$R2: \mathcal{LR}_{\mathsf{I;CA}} = 0$

	France	Germany	Italy	Japan	U.K.	U.S.
$\lambda_{I,CA,0}$	-0.20	0.30	-0.04	0.05	0.01	0.25
	(0.21)	(0.40)	(0.24)	(0.36)	(0.17)	(0.72)
$\lambda_{CA,I,0}$	-0.21	-0.61	-0.44	-0.21	-0.54	-0.15
	(0.17)	(0.24)	(0.20)	(0.23)	(0.17)	(0.09)
$\mathcal{LR}_{CA,I}$	-0.18	-0.50	-0.57	-0.20	-0.65	-0.28
	(0.18)	(0.21)	(0.19)	(0.26)	(0.12)	(0.22)
Wald Statistic	14.40	8.13	9.96	2.31	24.24	13.94
	[0.01]	[0.15]	[0.08]	[0.80]	[0.00]	[0.02]

In the top and bottom panels, standard errors appear in parenthesis and the brackets contain p-values. The Wald statistic and p-values in the top and bottom panel are based on the hypothesis $\lambda_{CA,I,j} = 0,\ j = 0,\ \ldots,\ 4$, and five degrees of freedom.

Table A4.2 Structural VARs

Sample Period: 1975.1 – 1995.4

$$R3: \lambda_{I;CA;0} = 0$$

	France	Germany	Italy	Japan	U.K.	U.S.
$\lambda_{CA,I,0}$	-0.34	-0.43	-0.47	-0.18	-0.53	-0.12
	(0.08)	(0.09)	(0.09)	(0.08)	(0.10)	(0.04)
$\mathcal{LR}_{I,CA}$	0.44	-0.24	0.05	-0.06	-0.02	-0.24
	(0.58)	(0.26)	(0.31)	(0.41)	(0.43)	(0.63)
$\mathcal{LR}_{CA,I}$	-0.24	-0.32	-0.59	-0.18	-0.64	-0.25
	(0.10)	(0.20)	(0.14)	(0.18)	(0.10)	(0.08)
Wald Statistic	29.69	23.80	31.50	6.21	43.12	22.30
	[0.00]	[0.00]	[0.00]	[0.29]	[0.00]	[0.00]

$$R4: \lambda_{CA;I;0} = 0$$

	France	Germany	Italy	Japan	U.K.	U.S.
$\lambda_{I,CA,0}$	-0.48	-0.49	-0.50	-0.29	-0.48	-0.96
	(0.12)	(0.10)	(0.10)	(0.13)	(0.09)	(0.28)
$\mathcal{LR}_{I,CA}$	-0.56	-0.69	-0.52	-0.39	-0.82	-1.04
	(0.43)	(0.23)	(0.21)	(0.40)	(0.19)	(0.53)
$\mathcal{LR}_{CA,I}$	-0.07	0.15	-0.21	-0.02	-0.44	-0.13
	(0.11)	(0.30)	(0.14)	(0.18)	(0.11)	(0.07)
Wald Statistic	11.05	1.29	3.72	1.37	10.46	9.52
	[0.03]	[0.86]	[0.44]	[0.85]	[0.03]	[0.05]

In the top and bottom panels, standard errors appear in parenthesis and the brackets contain p-values. The Wald statistic and p-values in the top (bottom) panel are based on the hypothesis $\lambda_{CA,I,j} = 0,\ j = 0, \ldots, 4$ ($j = 1, \ldots, 4$), and five (four) degrees of freedom.

Table A4.3 Structural VARs

Sample Period: 1975.1 – 1995.4

$$R5: \; \lambda_{\mathsf{CA;I;0}} \;=\; -1$$

	France	Germany	Italy	Japan	U.K.	U.S.
$\lambda_{I,CA,0}$	1.73	1.29	1.15	1.79	0.80	151.88
	(0.61)	(0.47)	(0.43)	(0.52)	(0.31)	(985.90)
$\mathcal{LR}_{I,CA}$	5.75	0.71	2.16	2.16	26.96	-8.37
	(6.35)	(0.74)	(2.20)	(1.85)	(254.35)	(5.93)
$\mathcal{LR}_{CA,I}$	-0.66	-0.82	-1.04	-0.83	-0.82	-0.93
	(0.17)	(0.20)	(0.20)	(0.31)	(0.10)	(0.42)
Wald Statistic	2.91	12.77	7.96	10.37	10.68	11.22
	[0.57]	[0.01]	[0.09]	[0.03]	[0.03]	[0.02]

$$R6: \; \mathcal{LR}_{\mathsf{CA;I}} \;=\; -1$$

	France	Germany	Italy	Japan	U.K.	U.S.
$\lambda_{I,CA,0}$	4.55	2.49	1.00	2.49	3.02	-132.86
	(4.87)	(2.55)	(0.76)	(2.06)	(3.80)	(1836.17)
$\lambda_{CA,I,0}$	-1.42	-1.26	-0.95	-1.22	-1.49	-1.10
	(0.38)	(0.37)	(0.21)	(0.54)	(0.41)	(0.79)
$\mathcal{LR}_{I,CA}$	34.23	1.38	1.78	3.09	-3.18	-7.48
	(192.75)	(1.78)	(2.30)	(3.97)	(2.14)	(7.08)
Wald Statistic	2.94	9.61	8.37	7.89	9.09	9.20
	[0.71]	[0.09]	[0.09]	[0.16]	[0.10]	[0.10]

In the top and bottom panels, standard errors appear in parenthesis and the brackets contain p-values. The Wald statistic and p-values in the top (bottom) panel are based on the hypothesis $\lambda_{CA,I,j} = 0,\ j = 1, \ldots, 4$ ($j = 0, \ldots, 4$), and four (five) degrees of freedom.

Table A5.1 FEVD under $R1: \mathcal{LR}_{CA,I} = 0$

Investment Response to the World Shock

Forecast Horizon	France	Germany	Italy	Japan	U.K.	U.S.
0	70.96	88.30	56.76	93.00	22.62	65.60
	(22.19)	(16.95)	(19.37)	(17.18)	(15.92)	(18.12)
2	70.89	88.06	56.76	93.00	17.42	65.40
	(22.36)	(17.25)	(19.37)	(17.19)	(15.11)	(18.22)
4	70.89	88.19	56.76	93.00	16.73	65.40
	(22.34)	(17.29)	(19.37)	(17.19)	(14.43)	(18.22)
12	71.09	88.33	56.76	93.00	16.87	65.31
	(22.25)	(17.33)	(19.37)	(17.19)	(13.68)	(18.25)
24	71.15	88.40	56.76	93.00	16.93	65.29
	(22.16)	(17.35)	(19.37)	(17.19)	(13.59)	(18.25)

Current Account Response to the Country-Specific Shock

Forecast Horizon	France	Germany	Italy	Japan	U.K.	U.S.
0	97.69	98.43	95.86	99.88	73.25	92.75
	(11.37)	(8.42)	(9.16)	(11.23)	(15.94)	(10.24)
2	97.88	98.11	95.86	99.88	75.30	93.34
	(11.27)	(8.69)	(9.16)	(11.23)	(16.32)	(9.86)
4	97.82	98.14	95.86	99.88	76.02	93.30
	(11.39)	(8.68)	(9.16)	(11.23)	(15.78)	(10.03)
12	97.71	98.12	95.86	99.88	77.74	93.36
	(11.07)	(8.70)	(9.16)	(11.23)	(14.02)	(10.06)
24	97.68	98.13	95.86	99.88	78.34	93.37
	(10.99)	(8.70)	(9.16)	(11.23)	(13.59)	(10.38)

In the top and bottom panels, small sample empirical standard errors appear in parenthesis. We generate 1000 replications of the SVAR to compute the empirical standard errors.

Table A5.2 FEVD under $R2: \mathcal{LR}_{I,CA} = 0$

Investment Response to the World Shock

Forecast Horizon	France	Germany	Italy	Japan	U.K.	U.S.
0	97.29	95.43	99.87	99.86	99.99	99.30
	(6.74)	(11.72)	(6.19)	(9.26)	(3.41)	(6.91)
2	97.27	95.54	99.87	99.86	99.60	99.32
	(6.75)	(11.81)	(6.19)	(9.26)	(3.72)	(6.89)
4	97.26	95.44	99.87	99.86	99.38	99.32
	(6.75)	(11.89)	(6.19)	(9.26)	(3.97)	(6.90)
12	97.30	95.37	99.87	99.86	98.15	99.33
	(6.69)	(11.76)	(6.19)	(9.27)	(4.20)	(6.90)
24	97.32	95.33	99.87	99.86	99.07	99.33
	(6.66)	(11.58)	(6.19)	(9.27)	(4.26)	(6.90)

Current Account Response to the Country-Specific Shock

Forecast Horizon	France	Germany	Italy	Japan	U.K.	U.S.
0	93.80	59.47	78.97	92.84	73.58	82.01
	(11.35)	(23.06)	(17.09)	(15.67)	(14.92)	(16.31)
2	93.53	58.20	78.97	92.84	69.45	81.69
	(11.74)	(23.39)	(17.09)	(15.67)	(16.81)	(16.73)
4	93.66	58.38	78.97	92.84	68.54	81.47
	(11.83)	(23.42)	(17.10)	(15.68)	(17.48)	(17.20)
12	93.91	58.29	78.97	92.84	67.80	81.29
	(11.90)	(23.54)	(17.10)	(15.68)	(17.90)	(17.51)
24	93.99	58.30	78.97	92.84	67.56	81.24
	(11.92)	(23.50)	(17.09)	(15.69)	(18.06)	(17.62)

In the top and bottom panels, small sample empirical standard errors appear in parenthesis. We generate 1000 replications of the SVAR to compute the empirical standard errors.

Table A5.3 FEVD under $R6$: $\mathcal{LR}_{CA,I}$ = -1

Investment Response to the World Shock

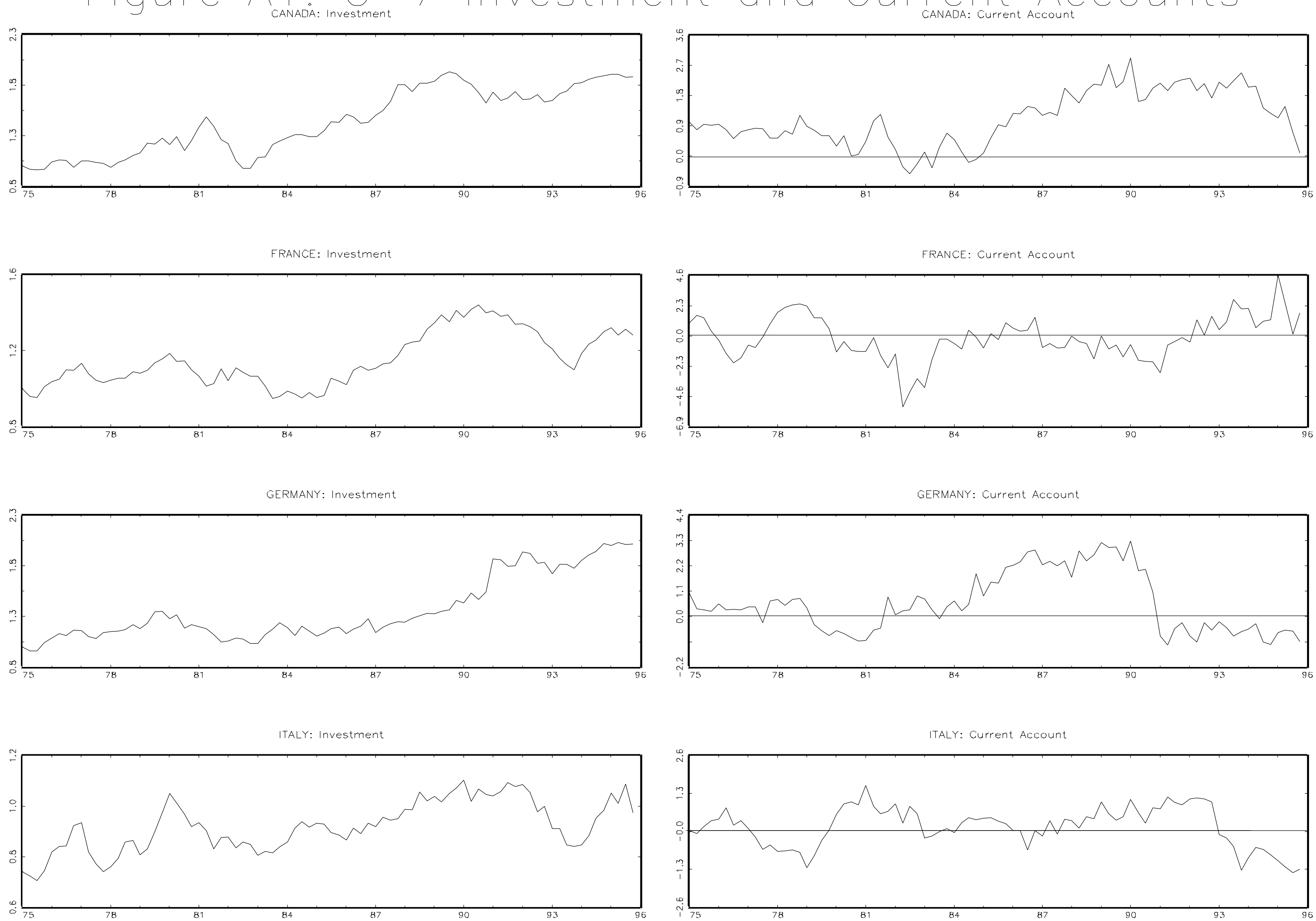
Forecast Horizon	France	Germany	Italy	Japan	U.K.	U.S.
0	34.49	49.92	75.45	36.28	47.58	10.78
	(17.18)	(21.96)	(18.83)	(23.62)	(20.99)	(15.96)
2	35.59	49.96	75.45	36.29	68.06	63.56
	(19.09)	(22.33)	(18.86)	(23.59)	(18.46)	(25.19)
4	37.40	43.37	75.45	36.28	86.30	100.00
	(22.84)	(23.25)	(18.50)	(24.04)	(16.91)	(34.85)
12	51.52	49.15	75.45	36.26	95.96	100.00
	(24.42)	(21.98)	(17.42)	(25.50)	(12.67)	(36.02)
24	85.50	49.06	75.45	36.25	94.51	100.00
	(22.61)	(20.70)	(16.40)	(25.75)	(11.06)	(36.62)

Current Account Response to the Country-Specific Shock

Forecast Horizon	France	Germany	Italy	Japan	U.K.	U.S.
0	4.38	9.30	26.42	16.09	31.07	00.04
	(9.42)	(14.89)	(18.35)	(19.62)	(12.96)	(8.12)
2	4.34	8.40	26.42	16.08	9.14	00.02
	(9.36)	(14.48)	(18.35)	(19.75)	(11.78)	(16.50)
4	5.15	8.57	26.41	16.06	6.61	0.00
	(10.03)	(14.80)	(18.35)	(20.02)	(11.99)	(23.32)
12	16.82	8.54	26.41	16.05	5.56	0.00
	(14.34)	(15.54)	(18.35)	(20.98)	(12.42)	(28.01)
24	6.02	8.54	26.41	16.05	5.52	0.00
	(15.34)	(15.63)	(18.32)	(21.04)	(12.83)	(28.51)

In the top and bottom panels, small sample empirical standard errors appear in parenthesis. We generate 1000 replications of the SVAR to compute the empirical standard errors.

Figure A1: G-7 Investment and Current Accounts



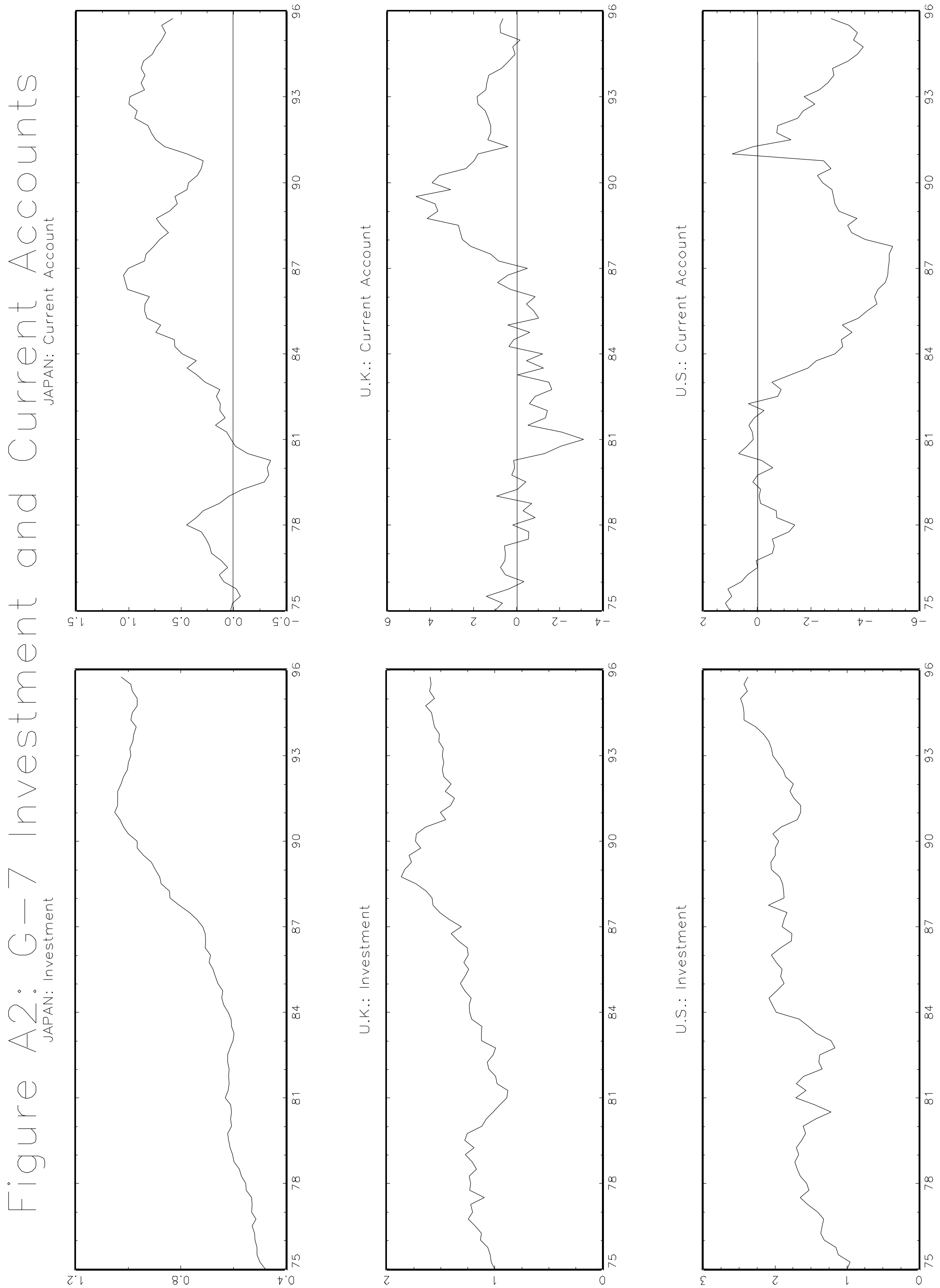


Figure A2: G-7 Investment and Current Accounts

Figure A3: French Results

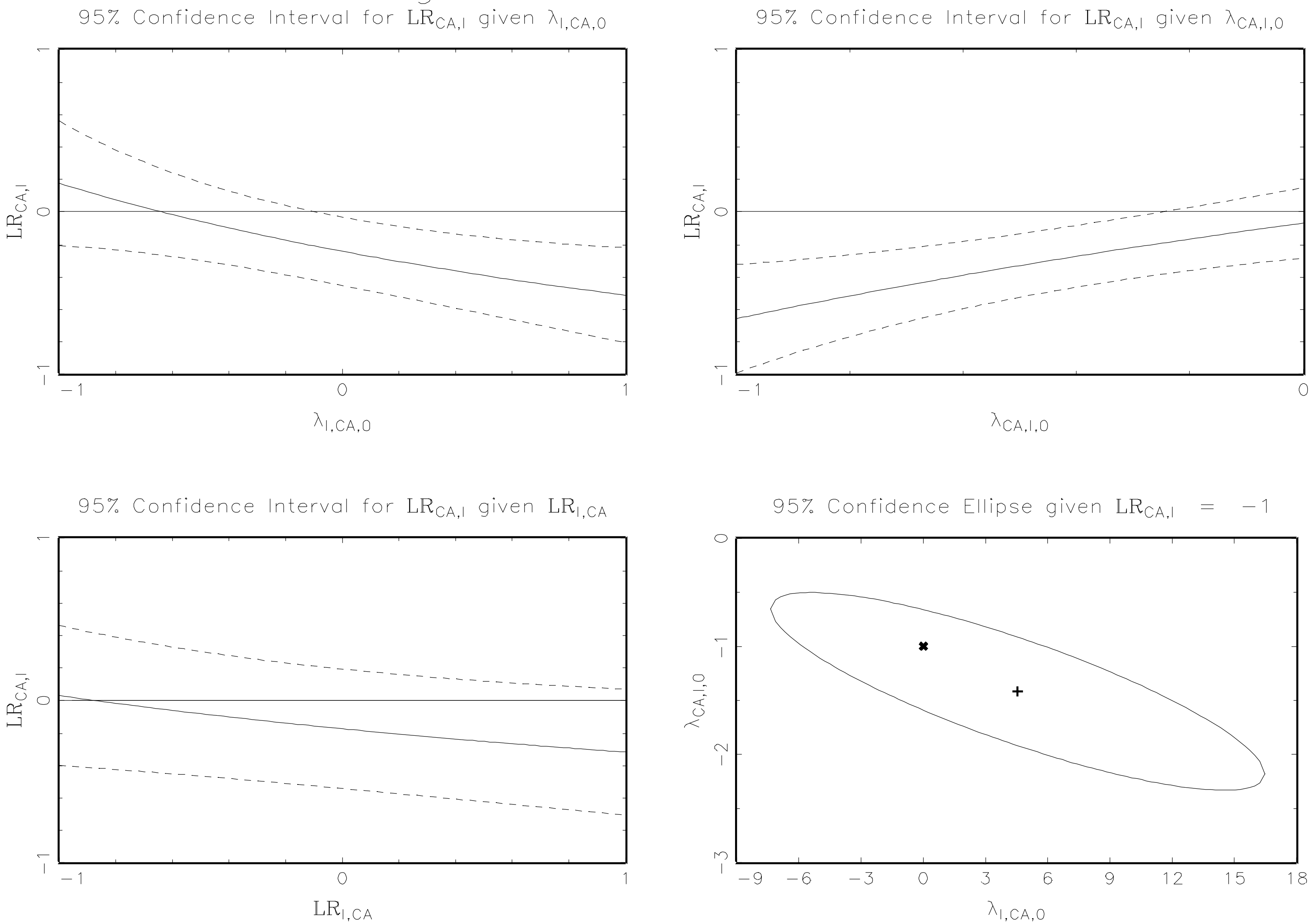


Figure A4: German Results

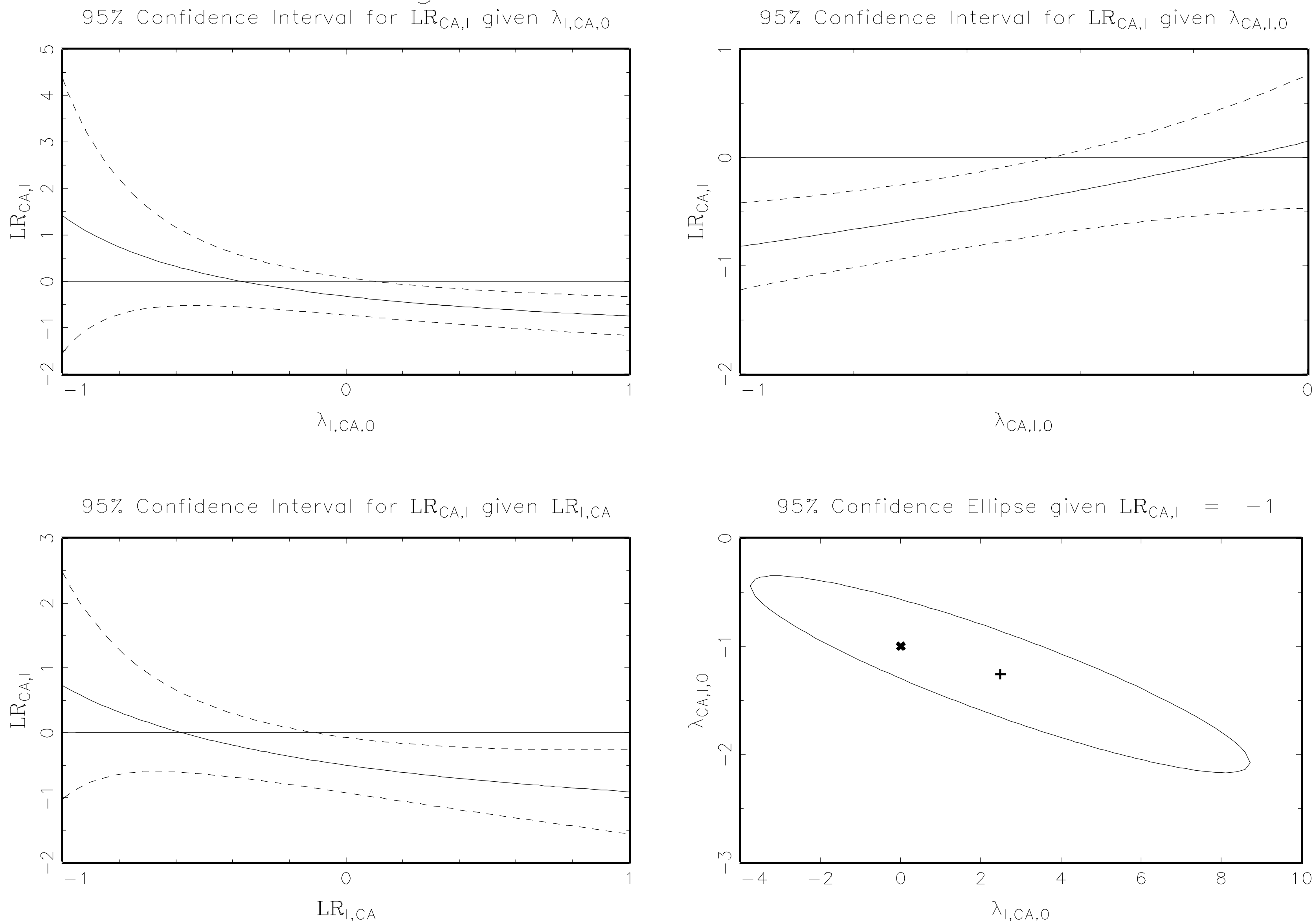


Figure A5: Italian Results

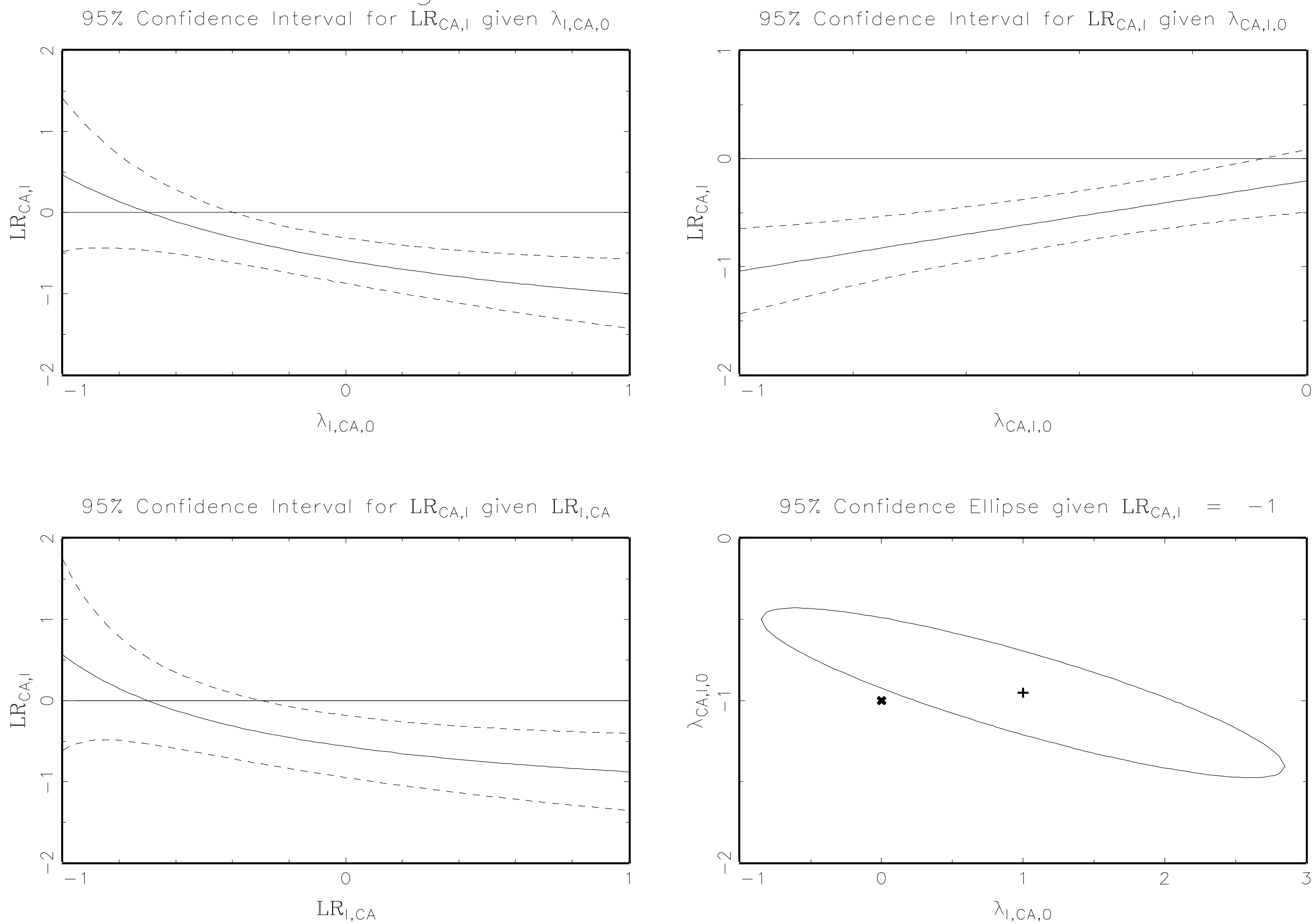


Figure A6: Japanese Results

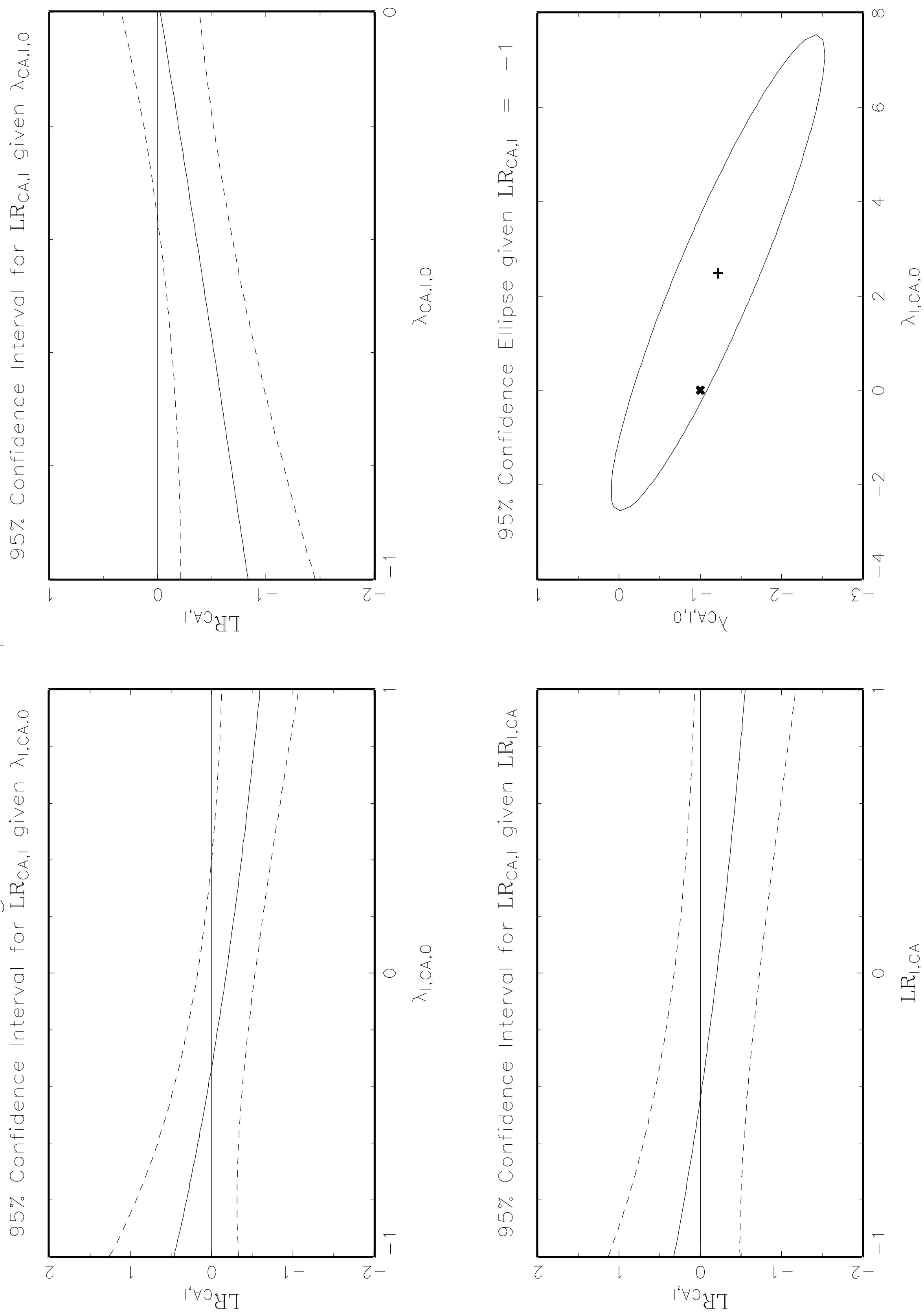


Figure A7: U.K. Results

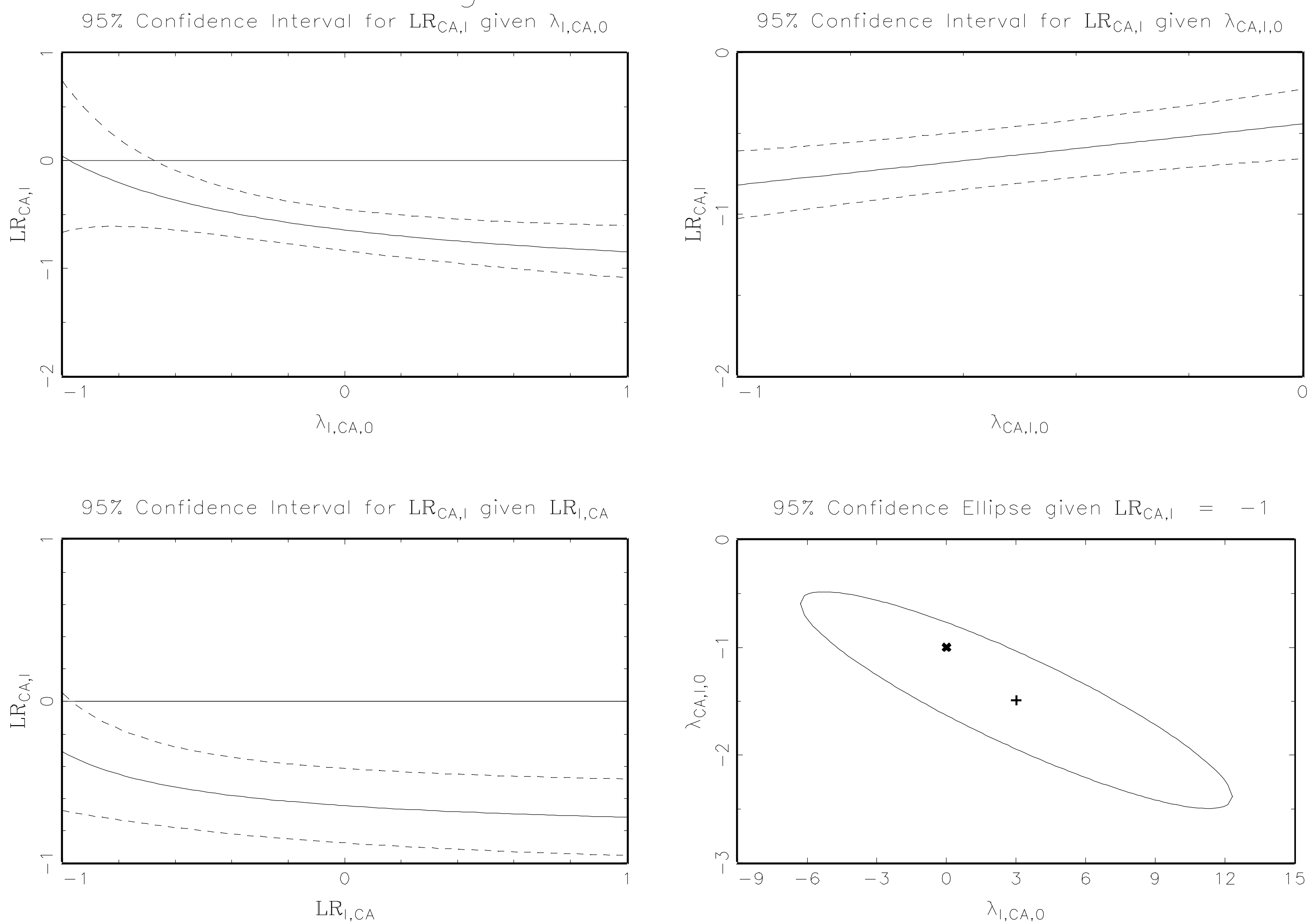


Figure A8: U.S. Results

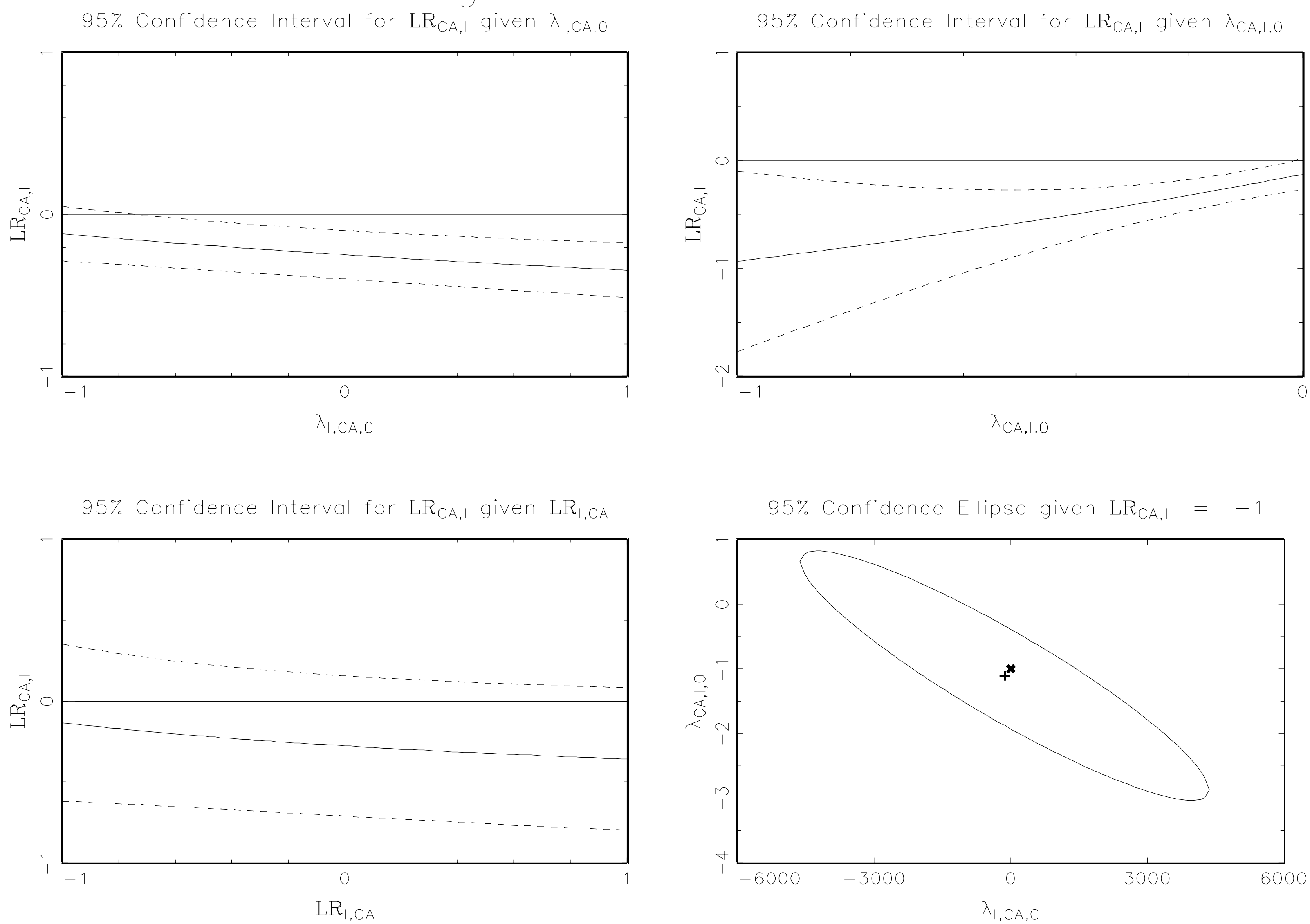


Figure A9: French 95% Confidence Ellipses

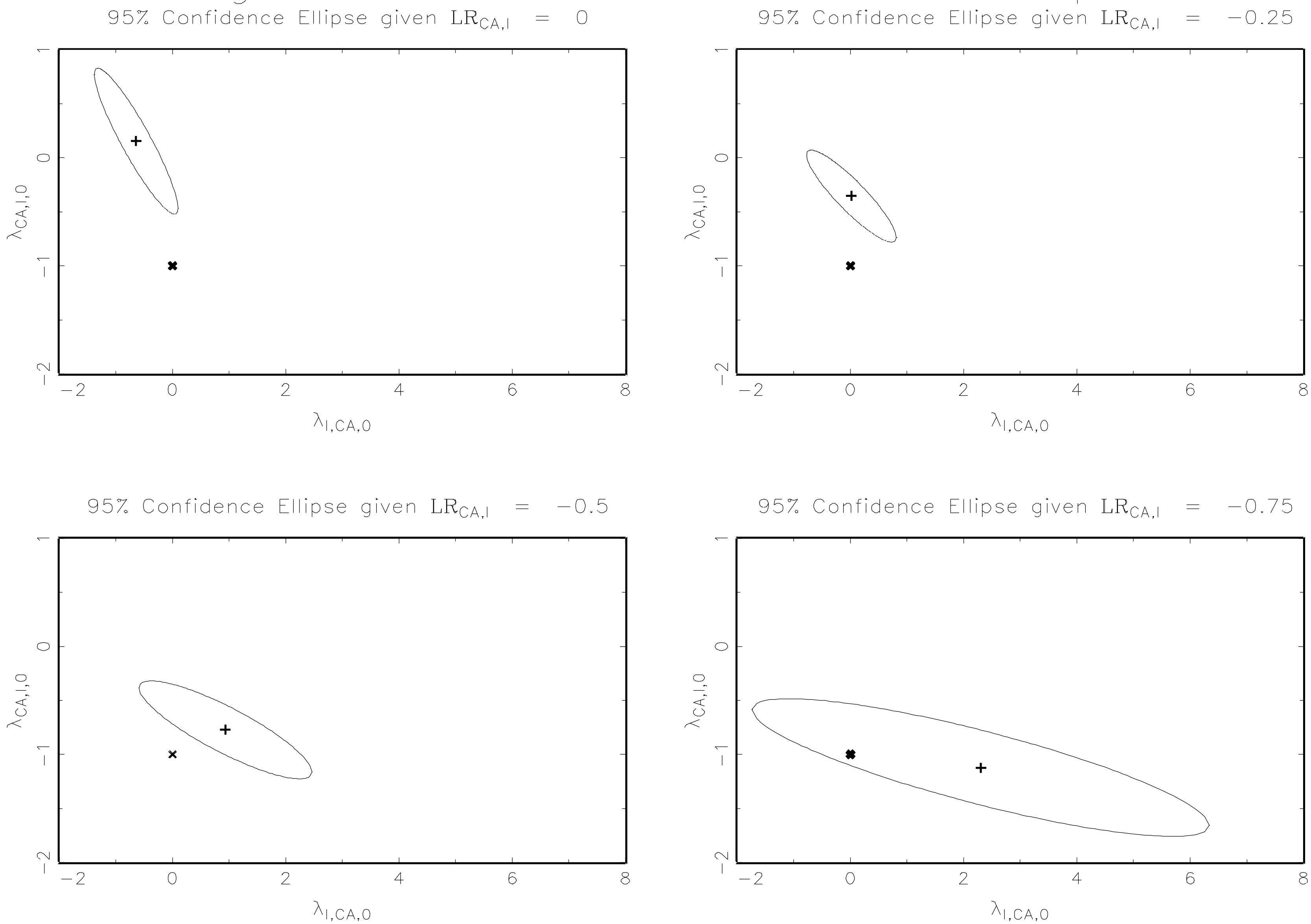


Figure A10: German 95% Confidence Ellipses

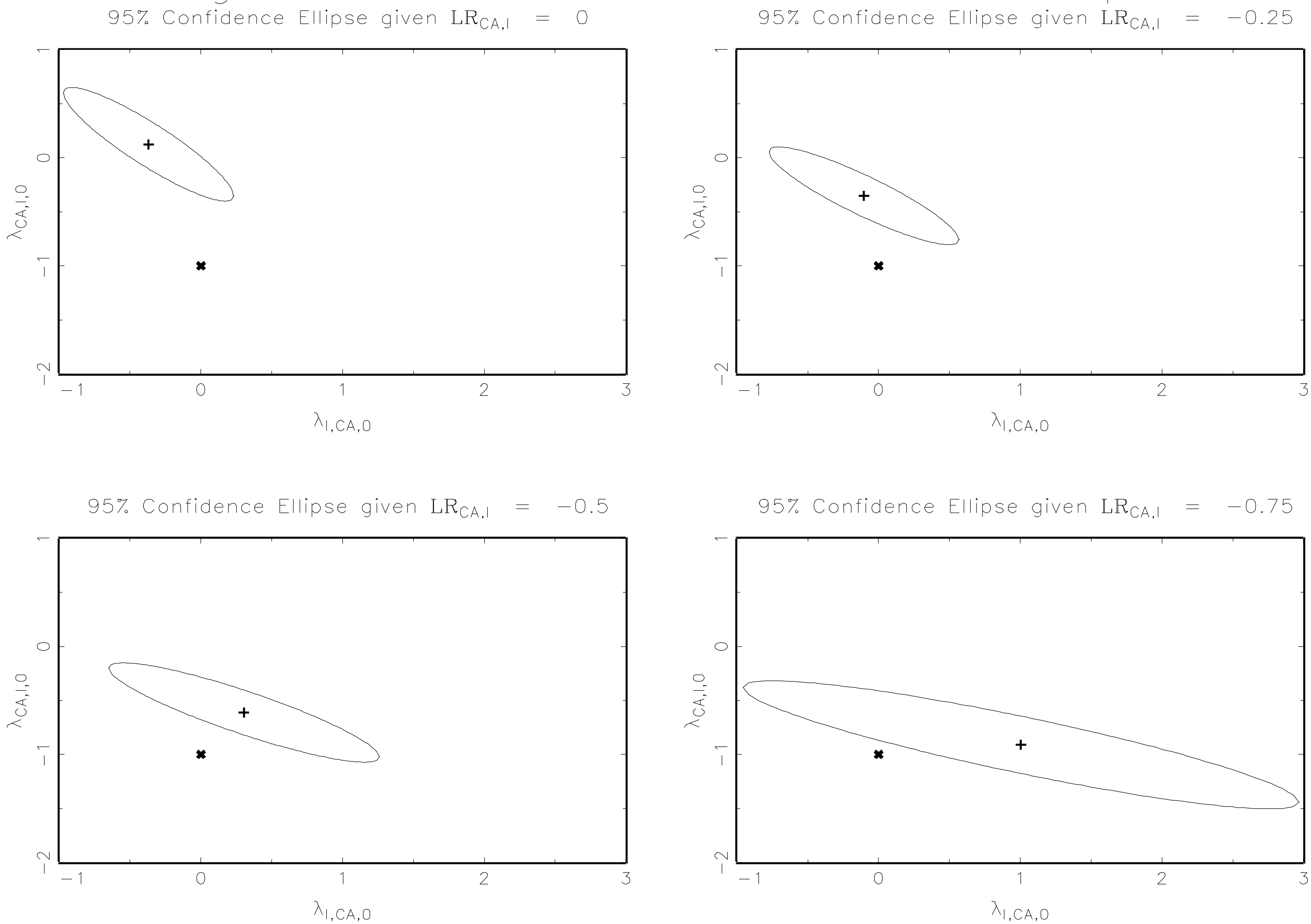


Figure A11: Italian 95% Confidence Ellipses

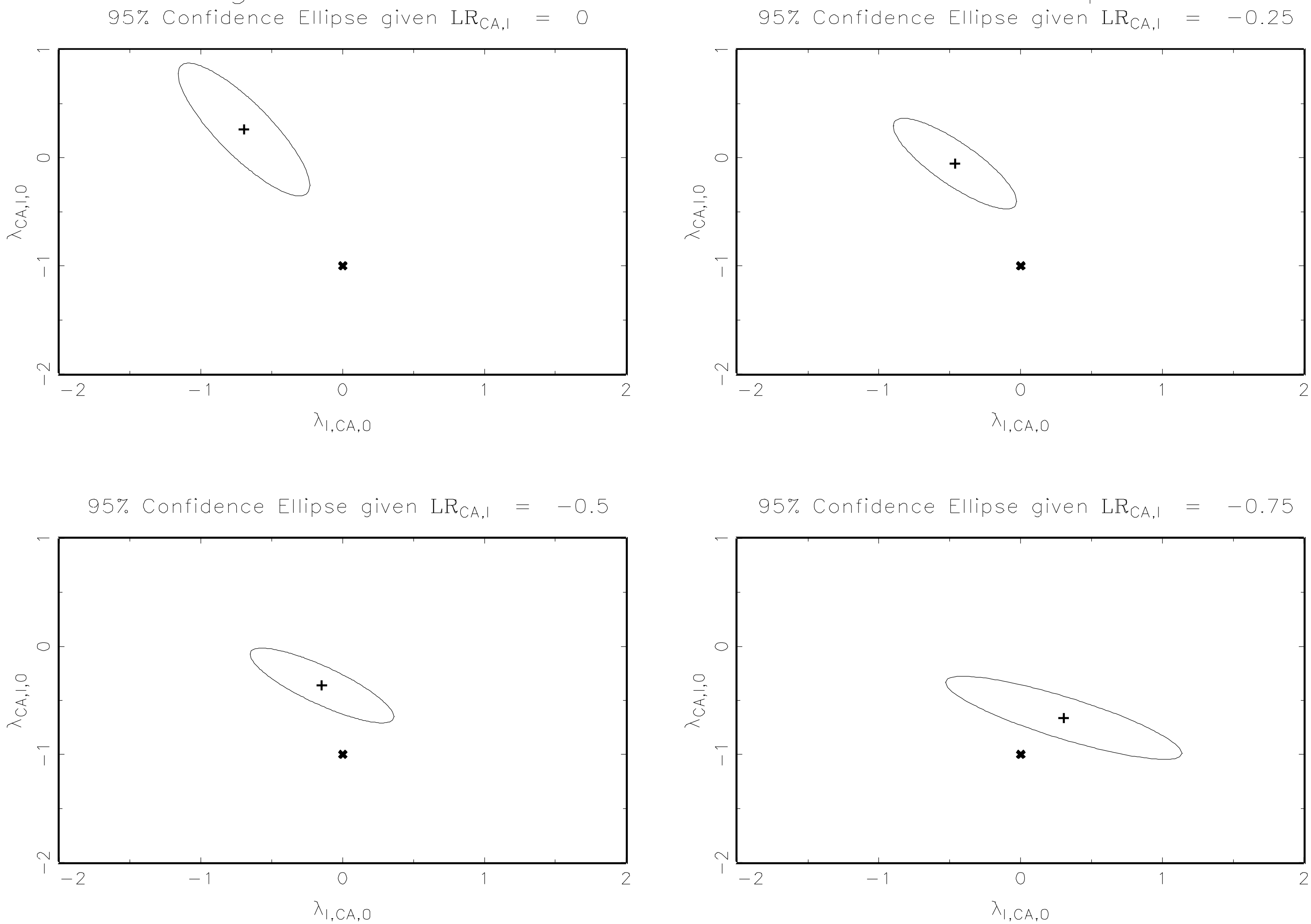


Figure A12: Japanese 95% Confidence Ellipses

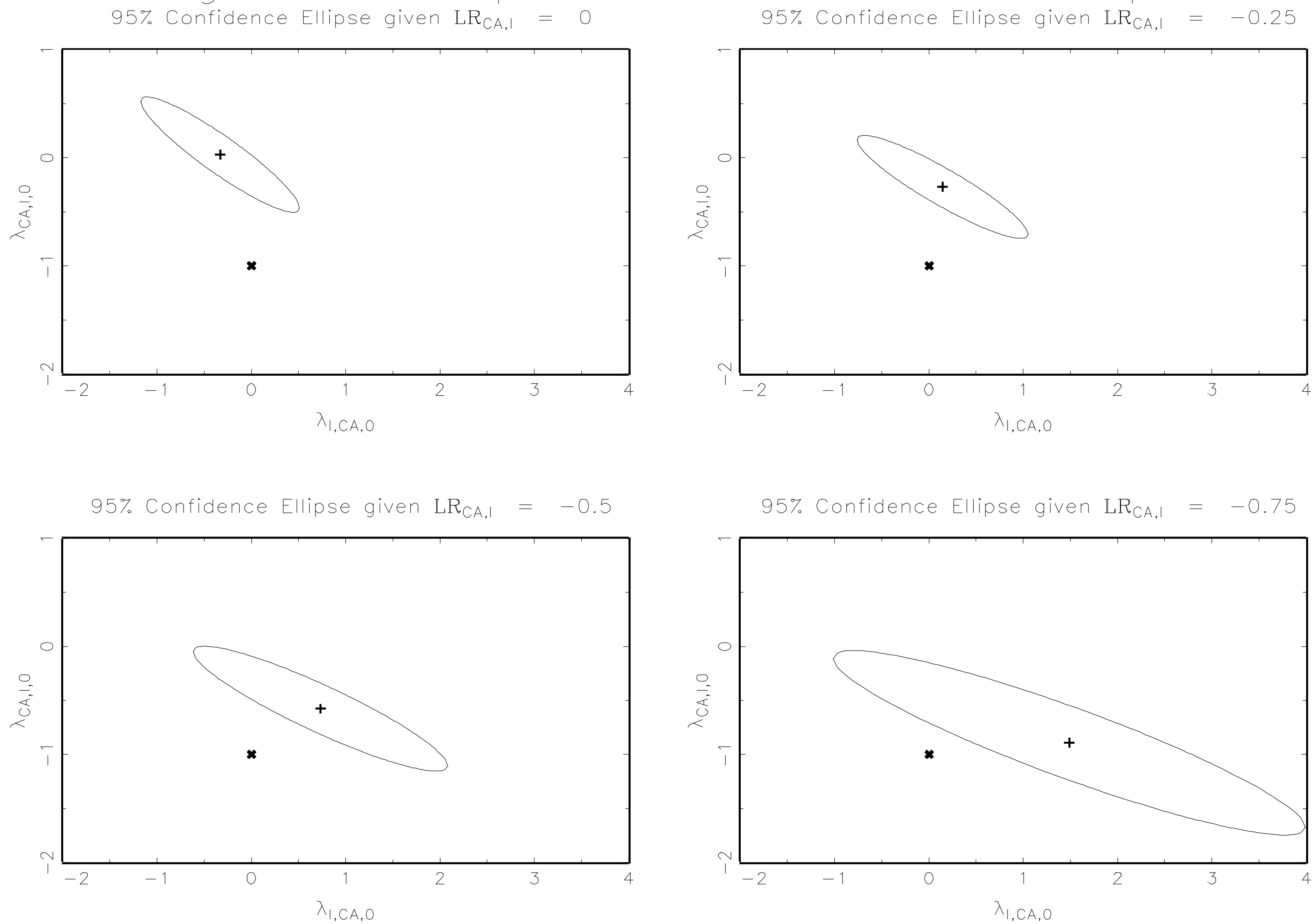


Figure A13: U.K. 95% Confidence Ellipses

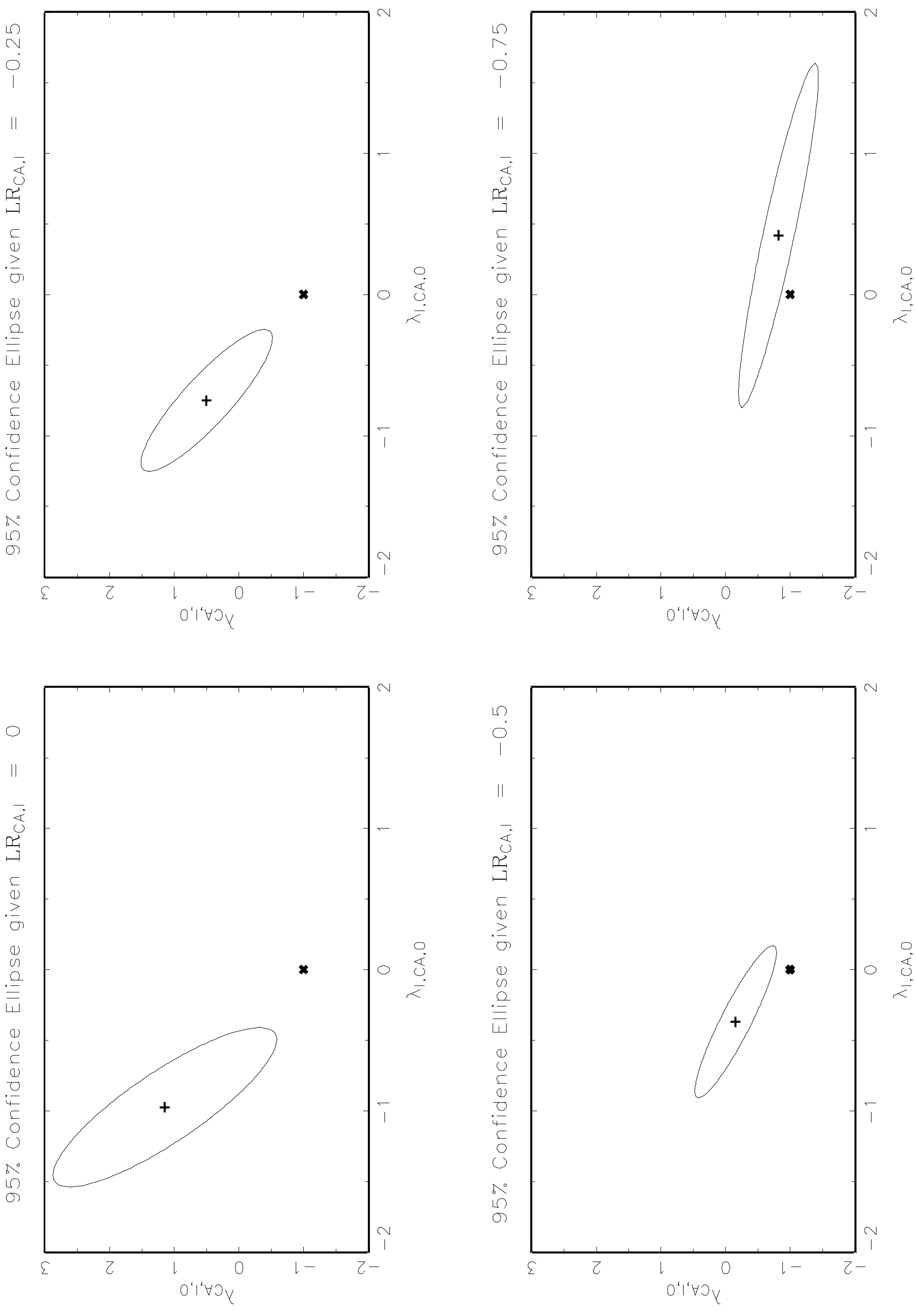


Figure A14: U.S. 95% Confidence Ellipses