Financial Regulation and Current Account Adjustment

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ABSTRACT

We study the effect of financial regulation on current account adjustment. The intertemporal model of the current account predicts that liqudity constraints, which we assume to partially reflect the degree of financial regulation, affect the size and persistence of the current account response to a domestic net output shock. This prediction is tested in a sample of 84 countries with an interacted panel VAR model where the coefficients are allowed to vary with the degree of financial regulation. Our results suggest that the reaction of the current account balance to a net output shock is 60% larger and substantially more persistent in a country with a low, than in one with a high, degree of financial regulation. This finding is robust to allowing the VAR coefficients to vary with other determinants of current account adjustment, such as the degree of capital account openess and the exchange rate regime.

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1. Introduction

The run-up to the global financial crisis of 2008 was characterized by large and persistent current account imbalances. One school of thought argues that 'the origins of the crisis lie in the imbalances in the world economy that built up over a decade or more' (King, 2009). There is both empirical and theoretical evidence to support this proposition. Theoretically, Caballero and Krishnamurthy (2009) show that capital flows can lead to a sharp rise in asset prices and a decrease in risk premia and interest rates. Empirically, Warnock and Warnock (2009) estimate that, in the absence of foreign official flows into US government bonds over the course of a year, long rates would be almost 100 basis points higher. Similarly, Sa, Towbin and Wieladek (2011) show that, in OECD countries, capital inflows have a greater effect on real house prices and residential investment than monetary policy shocks.

Others contend that the current account merely reflects domestic economic conditions. Laibson and Mollerstrom (2010) present a model in which a current account deficit is the result of a domestic asset price bubble and credit boom, features that the authors argue are consistent with the data. Similarly, Gete (2010) shows that increases in housing demand can lead to a trade deficit via consumption smoothing across tradable and non-tradable goods. Empirically, Fratzscher, Juvenal and Sarno (2010) find that US asset price movements can explain up to 32% of the US trade balance, providing support for this view.

A third, still nascent, school of thought points to financial deregulation/innovation as an explanation for global current account imbalances. In particular, Ferrero (2011) considers a shock to the collateral constraint, which he argues can be interpreted as the macroeconomic effect of financial innovation, in a two-country model. The realisation of this shock in a calibrated version of his model can, to some extent, replicate the US current account deficit and house price boom observed in the run up to the crisis.[1] In a similar vein, Borio and Disyatat (2011) argue that the international monetary and financial system suffers from excessive 'elasticity'[2], defined as 'the degree to which extent monetary and financial regimes constrain the credit creation process, and external funding more generally'. They hypothesize that an increase in the elasticity of the financial system over time, as a result of financial deregulation for instance, led to greater domestic and external imbalances. Indeed, the data suggests that the increase in the absolute size of current account imbalances (figure 1) and their persistence (figure 2) emerged against the backdrop of financial deregulation in both

[1] Favilukis, Ludvigson and Van Nieuwerburgh (2011) also explore the interaction of foreign flows and financial liberalisation but rule out household holdings of foreign assets by assumption.

[2] The authors trace this idea back to as early as Jevons (1875).

OECD and emerging market countries (figure 3)[3], providing informal support for the 'excess elasticity' hypothesis. But, to our knowledge, no previous work has rigorously tested if, and to which extent, current account adjustment depends on the domestic financial regulatory regime. In this paper, we take a step towards filling this gap. In particular, we provide the first empirical evidence that, for a given set of net output shocks, the current account response is larger and more persistent in a country with a low, than in one with a high, degree of financial regulation.

In the first part of the paper, we derive robust identification restrictions and theoretical predictions from the intertemporal model of the current account (Sachs, 1981), which we augment with liquidity constraints. Following previous theoretical and empirical work on long-run economic growth, savings, and consumption, these constraints are assumed to partially reflect the degree of financial reguation[4]. Recent work argues that the standard model needs be extended by either external habit formation (Bussiere, Fratzscher and Mueller, 2008), internal habit formation (Gruber, 2004) or a stochastic world real interest rate (Bergin and Sheffrin, 2000) to match the persistence of the current account typically observed in the data. This theory suggests that a net output shock will have a larger effect on the current account balance if agents are less liquidity constrained, a prediction that we show is robust to all of these assumptions. In the case of external habit formation in consumption, the current account response also becomes more persistent as the liquidity constraint eases. In other words, the lighter financial regulation, the greater and more persistent the reaction of the current account to net output shocks. The intuition is simple: the current account reflects agents' savings decisions in response to net output shocks. In a repressed (liberalized) financial system, few (many) agents have access to borrowing and saving and the current account will therefore show a smaller (greater) reaction to domestic shocks.

In the second part of the paper, we take our theory to the data. The theoretical model implies a VAR data generating process that consists of real per capita net output growth, the fiscal balance to net output ratio and the current account to net output ratio. These theoretical VAR coefficients, that is the effect of a net output shock on the current account

[3]We define persistence as the AR(1) coefficient from an autoregressive panel data model, estimated with country fixed effects, of the current account to GDP ratio.

[4]We do not make an explicit distinction between financial regulation and financial repression. While the aim of financial regulation (repression) is financial stability (cheap government financing), the policies used to achieve these objectives are not mutually exclusive. Even the minimum capital requirement, intended to promote financial stability, incentivise banks to hold a larger fraction of government securities, as a result their zero risk-weight, than they would have otherwise. In this sense, the distinction between financial regulation and repression, at least at our level of aggregation, is one of semantics, rather than substance.

balance, depend on the degree of liquidity constraints (financial regulation). We therefore allow the structural VAR coefficients of our empirical model to vary with the degree of financial regulation and capital account openness, an independent determinant of liquidity constraints (Lewis, 1997). This interacted panel VAR model is estimated with time and fixed effects on an unbalanced panel of annual data across 84 countries over the period 1973-2005, using the *de jure* financial regulation indices provided in Abiad, Detragiache and Tressel (2010). Finally, the theory also provides identification restrictions which are robust to parametrization and therefore naturally translate into sign and shape restrictions of the type first introduced in Canova and De Nicolo (2002), Faust and Rogers (2003) and Uhlig (2005). We use these restrictions to identify a log level (difference) net output shock and then compare the associated current account to net output impulse responses between regimes of low and high financial regulation. Our results confirm that the impact of both shocks on the current account differs across these two regimes in a statistically-significant way. The response of the current account to either type of net output shock is approximately 60% larger, as well as more persistent, in the average financially-liberalized country than in the average financially-repressed country. This implies that the degree of financial regulation is an important determinant of current account adjustment. To our knowledge, this insight is not part of the current monetary and financial reform debate, but could have potentially important implications for the adjustment of global imbalances going forward.

This paper contributes to the existing literature in a number of ways. We provide the first empirical investigation of the interaction between financial regulation and the current account, expanding upon existing empirical work that examines the relationship between financial regulation and savings decisions in the closed economy (see Bayoumi (1993b), Bayoumi and Koujiano (1989), Japelli and Pagano (1994), and Bandiera et al (2000) among others.) Second, we work with a broad sample of countries for the standards of previous papers studying the effect of financial regulation on savings decisions, being the first to apply the sign restriction methodology to test the predictions of this type of model. Finally, a minor contribution is our detailed exploration of the role of liquidity constraints in several standard versions of the intertemporal model of the current account.

The rest of this paper proceeds as follows. Section 2 lays out the theoretical model and derives theoretically robust identification restrictions and predictions. Section 3 discusses the empirical specification and the data used. Section 4 presents the results and robustness checks. Section 5 concludes.

2. Theory

Our aim is to embed financial regulation in a model of the current account that is simple enough to take to the data in a neat way. We build on the Sachs (1981) intertemporal model of the current account (ICA). Conceptually, financial regulation can affect macroeconomic variables in at least two ways. First, regulation can affect the efficiency of transforming savings into productive investment (Goldsmith, 1969) by promoting or hindering competition in the banking system. Second, the volume of savings flowing into investment can either increase (Mckinnon, 1973; Shaw, 1973) or decrease (Devereux and Smith, 1994; Japelli and Pagano, 1994) following deregulation. The former is likely to affect the price, while the latter the quantity, of capital available for investment. Since current account balances reflect quantities rather than prices we focus on the second channel. While theoretically ambiguous, empirical evidence supports the proposition that tighter financial regulation decreases savings through the liquidity constraints channel. Diaz-Alejandro (1985) points out that in developing countries, the volume of savings tends to fall following financial liberalization (deregulation). Bayoumi (1993a), Bayoumi (1993b), and Sarno and Taylor (1998) find that financial deregulation in the UK in the 1980s decreased liquidity constraints, leading to a decline in aggregate savings. Bayoumi and Koujianou (1989) and Japelli and Pagano (1994) confirm this pattern across a range of industrial countries. Bandiera, Caprio, Hohonan and Schiantarelli (2000) also find empirical support for the idea that financial regulation decreases savings via the liquidity constraints channel in eight emerging market economies. Finally, using data for 72 countries, Lewis (1997) finds that consumers in countries with government restrictions on international transactions tend to act as if they are liquidity constrained.[5] This breadth of evidence justifies our approach of introducing financial regulation via liquidity constraints in the ICA model.

Bergin and Sheffrin (2000) and Ghosh (1995) among others, do not find empirical support for the simple intertemporal current account model in most G7 countries. The literature has proposed three different modifications of the basic model to improve its empirical performance. Gruber (2004) introduced internal habit formation (where utility is a function of past individual consumption), Bussiere, Fratzscher and Mueller (2008) used external habit formation (where utility is a function of past average consumption), and Nason and Rogers (2006) argued that a time-varying stochastic world interest rate delivers a more realistic model. Recently, Kano (2009) has shown that the internal habits and time-varying world real interest rate approaches provide observationally equivalent predictions. Together with

[5] Note that in contrast to her work, we focus on government restrictions on domestic, rather than international, financial transactions.

one of these three frictions, we add liquidity constraints to the standard ICA model in order to capture the effects of financial regulation. This allows us to show that our identification restrictions are robust across these modeling choices and reasonable parametrization.

We consider a small open, endowment economy populated by a large number of households who maximize utility subject to a budget constraint. Output, investment, government expenditure, and lump-sum taxes are exogenous.[6] There are two types of households in the economy. One type of households is liquidity constrained in the sense that they do not have access to any savings technology. We will refer to them as non-Ricardian households. These households consume their exogenous income each period:

$$C_t^{NR} = Y_t - I_t - T_t$$

where Y_t is output, I_t is investment and T_t are taxes. The second type of household has access to incomplete international financial markets in which only non-contigent riskless bonds are traded. These households are characterized by their optimal behavior with respect to the intertemporal allocation of consumption. We will refer to this group of households as Ricardian throughout. We assume that non-Ricardian households make up a fraction of $\lambda \epsilon [0,1]$ of the population with Ricardian households as the remainder. Hence, aggregate consumption is defined as $C_t = \lambda C_t^{NR} + (1-\lambda)C_t^R$.

The representative Ricardian household solves the following maximization problem:

$$U_t = E_0 \sum_{t=0}^{\infty} \beta^t U(C_t^R - hC_{t-1}) \tag{1}$$

$$s.t. B_{t+1}^P = (1+r_t)B_t^P + Y_t - I_t - T_t - C_t^R \tag{2}$$

$$\underset{i \to \infty}{Lim} R_{t,i} E_t B_{i+1} \geq 0 \tag{3}$$

$$R_{i,t} = \left\{ \begin{matrix} 1/(\prod_{j=t+1}^{t+i} (1+r_j)) & \text{if } i \geq 1 \\ 1 & \text{if } i = 0 \end{matrix} \right\}$$

where $\beta \in \{0,1\}$ is the discount factor and C_t^R consumption at time t, and C_{t-1} is past consumption. B_{t+1}^P is the net stock of international bonds held by the Ricardian agent at the

[6]The exogeneity assumption is a convenient simplification that is unlikely to hold empirically. However, since our main theoretical predictions follow directly from the Euler equation, they would most likely also hold in more complex DSGE models where all macroeconomic variables are endogenous. Unlike in those models however, the fact that we can obtain an analytical solution and we do not need to resort to numerical methods, allows us to easily and transparently demonstrate that our predictions are robust to reasonable parameter values.

end of time t, which are reimbursed at the world real interest rate r_t. Households choose C_t^R and B_{t+1}^P to maximize discounted lifetime consumption subject to the budget contraint and a no-ponzi condition. We assume $0 \leq h < 1$, which implies that Ricardian households may be habit forming with respect to consumption. We explore both the possibility of external and internal habit formation in consumption. As a result, utility is increasing in consumption expenditure that exceeds the depreciated value of last period's average consumption in the economy (in the case of external habits) or the depreciated value of the household's lagged consumption, with h being the rate of depreciation. The utility function takes a log form throughout. The first-order necessary conditions of this optimization problem comprise the budget constraint (2), the tranversality condition (3) and the following Euler equation:

$$U'(C_t^R - hC_{t-1}) = \beta E_t[(1 + r_{t+1})U'(C_{t+1}^R - hC_t)] \tag{4}$$

Iterating the budget constraint and imposing the no-ponzi condition yields the intertemporal budget constraint of the Ricardian agent:

$$E_t \sum_{i=0}^{\infty} R_{t,i} C_{t+i}^R = (1 + r_t)B_t^P + E_t \sum_{i=0}^{\infty} R_{t,i}(Y_{t+i} - I_{t+i} - T_{t+i}) \tag{5}$$

We also assume the presence of an exogenous government sector. The government budget constraint is:

$$B_{t+1}^G = (1 + r_t)B_t^G + T_t - G_t$$

where B_{t+1}^G is the net stock of international bonds held by the government at the end of time t. G_t is government spending. Iterating the government budget constraint and combining it with (5), one can show that

$$E_t \sum_{i=0}^{\infty} R_{t,i} C_{t+i}^R = (1 + r_t)(B_t^P + B_t^G) + E_t \sum_{i=0}^{\infty} R_{t,i} NO_{t+i} \tag{6}$$

where $NO_t = Y_t - I_t - G_t$ is net output. Since only Ricardian agents hold bonds, the stock of external assets, B_t, is

$$B_t = (1 - \lambda)B_t^P + B_t^G$$

In order to derive an analytical solution to the present value of the current account, we take linear approximations of both the budget constraint (6) and the Euler equation (4). Following the approach in Kano (2008), one can show after a fair amount of algebra (see appendix) that

$$\begin{aligned} & \frac{C_t^R}{NO_t}[1+\sum_{i=1}^{\infty} E_t \exp\left\{\sum_{j=t+1}^{t+i} (\Delta \ln C_j^R - \ln(1+r_j))\right\}] \\ = \;& \exp\{\ln(1+r_t)\} \frac{B_t^P + B_t^G}{NO_t} + [1+\sum_{i=1}^{\infty} E_t \exp\left\{\sum_{j=t+1}^{t+i} (\Delta \ln NO_j - \ln(1+r_j))\right\}] \end{aligned} \tag{7}$$

Let c, b^G, b^P, μ and γ denote the unconditional means of the consumption to net output ratio, the net foreign private asset to net output ratio, the net foreign public asset to net output ratio, the log of the gross world real interest rate and the growth rate of net output and consumption, respectively. Let us define $\widetilde{X_t} = X_t - \overline{X}$, where $\overline{X}$ is the steady state value of variable X. Taking a first-order Taylor expansion of (7) around the steady state yields:

$$\begin{aligned} \widetilde{\frac{C_t^R}{NO_t}} \;=\;& \frac{1-\kappa}{e^{-\mu}}(\widetilde{\frac{B_t^G}{NO_t}} + \widetilde{\frac{B_t^P}{NO_t}}) + \sum_{i=1}^{\infty} \kappa^i E_t \left\{\Delta\widetilde{\ln NO}_{t+i} - \ln\widetilde{(1+r_{t+i})}\right\} \\ & -c \sum_{i=1}^{\infty} \kappa^i E_t \left\{\Delta\widetilde{\ln C^R}_{t+i} - \ln\widetilde{(1+r_{t+i})}\right\} + \frac{1-\kappa}{e^{-\mu}}(b^G + b^P) \ln\widetilde{(1+r_t)} \end{aligned} \tag{8}$$

where $\kappa = \exp(\gamma - \mu)$. This is the linearized budget constraint. At this point further assumptions have to be made. The literature has proposed three different simplifications to solve the model. The first option is to assume that the world real interest rate is constant and external habit formation (i.e., the argument of the utility function is $C_t^R - hC_{t-1}$ where C denotes average consumption in the economy). The second option is to assume a constant world real interest rate and internal habit formation (i.e., the argument of the utility function is $C_t^R - hC_{t-1}^R$). Finally, one can assume a stochastic world real interest rate and the absence of habit formation ($h = 0$ in our model). Kano (2009) shows that these last two frictions imply observationally equivalent data generating processes. Since in practice it is difficult to know which model is underlying the data generating process, we show that our identification assumptions are robust to any of these specific modeling choices. The detailed derivation of the current account reaction functions is relegated to the appendix.

Under a constant world real interest rate and external habit formation the current account reaction function is

$$\begin{aligned}\widetilde{ca}_t &= (1-\lambda)h\widetilde{ca}_{t-1} + \lambda\widetilde{fb}_t - (1-h\kappa)(1-\lambda)\sum_{i=0}^{\infty}\kappa^i E_t\left\{\Delta\widetilde{\ln NO}_{t+i}\right\} \\ &\quad +(1-\lambda)\Delta\widetilde{\ln NO}_t - h\kappa(1-\lambda)^2\sum_{i=0}^{\infty}\kappa^i(E_t-E_{t-1})\left\{\Delta\widetilde{\ln NO}_{t+i}\right\} + f_t\end{aligned} \tag{9}$$

where f_t is an expectational error, defined as $f_t = h(1-\lambda)^2 c\sum_{i=1}^{\infty}\kappa^i(E_t-E_{t-1})\left\{\Delta\widetilde{\ln C^R}_{t+i-1}\right\}$, $\widetilde{ca}_t = \widetilde{\frac{CA_t}{NO_t}}$ and $\widetilde{fb}_t = r\widetilde{\frac{B_t^G}{NO_t}} + \widetilde{\frac{T_t}{NO_t}} - \widetilde{\frac{G_t}{NO_t}}$. Let us develop some intuition underlying this equation. In the absence of liqudity constraints, the current account under external habit formation is a function of the lagged current account and a weighted average of current and future net output changes.[7] As habits become stronger, the importance of both the lagged current account to net output ratio increases (the first term in (9)) and the weight of expected future net output growth (third term) increases. Liquidity constraints play three roles in equation (9). First, they create an interaction between the fiscal balance and the current account. This is because the way government spending is financed directly affects the income and consumption of non-Ricadian households. A fiscal deficit results in a current account deficit since it increases the income of these households.[8] Second, as a larger fraction of households is liquidity constrained, the importance of net output shocks diminishes since fewer households are able to smooth such shocks by borrowing and lending internationally. Finally, the coefficient on the lagged current account term is a function of both habits and the fraction of liquidity constrained consumers. This follows directly from the Euler equation, since agents desired smoothing pattern depends on aggregate average consumption and hence implicitly on the liquidity constraint. With external habits, therefore, the speed of current account adjustment also becomes a function of the liqudity constraint.

Under a constant world real interest rate and internal habitual consumption the current account reaction function is

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[7] In this case the model becomes virtually identical to Gruber (2004).

[8] This is consistent with empirical evidence that documents twin deficits in the data (Kim and Roubini, 2008 and Corsetti and Mueller, 2008).

$$
\begin{aligned}
\widetilde{ca}_t &= h\widetilde{ca}_{t-1} + \lambda\widetilde{fb}_t - \lambda h\widetilde{fb}_{t-1} + (1-\lambda)\Delta\widetilde{\ln NO}_t \\
&\quad -(1-\lambda)(1-h\kappa)\sum_{i=0}^{\infty}\kappa^i E_t\left\{\Delta\widetilde{\ln NO}_{t+i}\right\} \\
&\quad -(1-\lambda)(h\kappa)\sum_{i=0}^{\infty}\kappa^i (E_t - E_{t-1})\left\{\Delta\widetilde{\ln NO}_{t+i}\right\} + f_t
\end{aligned} \tag{10}
$$

where f_t is an expectational error, defined as $f_t = (1-\lambda)\frac{c\kappa h}{1-\kappa h}(\Delta\widetilde{\ln C^R}_t - E_{t-1}\Delta\widetilde{\ln C^R}_t)$ and $\widetilde{ca}_t = \widetilde{\frac{CA_t}{NO_t}}$. The intuition underlying (9) and (10) is very similar. Unlike in (9), the lagged current account coefficient is not affected by liquidity constraints in (10). This is because, in the case of internal habitual consumption, only individual consumption is relevant and there is no scope for the liquidity constraint to interact with the lagged current account term.

Without habit formation, but under the assumption of a stochastic time-varying world real interest rate ($\widetilde{r}_t = \rho^r \widetilde{r}_{t-1} + \eta_t$), the current account reaction function is

$$
\begin{aligned}
\widetilde{ca}_t &= \rho^r\widetilde{ca}_t + \lambda\widetilde{fb}_t - \rho^r\lambda\widetilde{fb}_{t-1} + (1-\lambda)\Delta\widetilde{\ln NO}_t \\
&\quad -(1-\lambda)(1-\rho^r\kappa)\sum_{i=0}^{\infty}\kappa^i E_t\left\{\Delta\widetilde{\ln NO}_{t+i}\right\} \\
&\quad -(1-\lambda)(\rho^r\kappa)\sum_{i=0}^{\infty}\kappa^i (E_t - E_{t-1})\left\{\Delta\widetilde{\ln NO}_{t+i}\right\} + f_t
\end{aligned} \tag{11}
$$

where $f_t = [((1-\lambda)(e^{\mu}-1)b^G - e^{\mu}b^G) + \frac{(1-\lambda)\rho^r\kappa}{1-\rho^r\kappa}]\eta_t$. Note that, as in Kano (2009), the current account reaction function under internal habits (equation 10) is observationally equivalent to the reaction function under a stochastic world real interest rate (equation 11). For the remainder of this paper, we will therefore only derive results for the internal habits model, keeping in mind that its prediction are observationally equivalent to the time-varying stochastic world real interest rate model.

2.1. How does financial regulation affect the response of the current account to output shocks?

In this section we derive robust identification restrictions and examine how the current account reacts to net output shocks at different levels of regulation (the liquidity constraint) in our theoretical model. In order to understand how the current account reacts to net output shocks at different levels of financial regulation, we need to make further assumptions on the stochastic process driving $\widetilde{\ln NO_t}$. Net output can be subject to shocks in log differences, $\Delta \ln NO_t = \rho \Delta \ln NO_{t-1} + \varepsilon_t$, or log levels, $\ln NO_t = \rho \ln NO_{t-1} + \varepsilon_t$. A priori, it is not feasible to know which process is driving the log of net output. We will therefore consider both of them and examine the effect of an unexpected shock to either net output process on the current account under the assumption of internal and external habits individually. At this point, most previous work would solve the model numerically and show that those theoretical impulse responses, from which the sign identification restrictions are derived, are robust to many different possible parameter values.[9] The advantage of our approach is that we can demonstrate the robustness of our identification restrictions and theoretical predictions analytically.

Let us consider the case of external habitual consumption first. Under the assumption that net output is subject to shocks in log differences, one can show that (see appendix) (9) becomes

$$\begin{aligned} \widetilde{ca}_t &= (1-\lambda)h\widetilde{ca}_{t-1} + \lambda\widetilde{fb}_t + \frac{(1-\lambda)\kappa(h-\rho)}{1-\rho\kappa}\rho\Delta\widetilde{\ln NO}_{t-1} \\ &\quad -\kappa(1-\lambda)\frac{(\rho(1-\kappa)+h(\kappa-\lambda)-\rho\kappa h(1-\lambda))}{(1-\kappa)(1-\rho\kappa)}\varepsilon_t + f_t \end{aligned} \tag{12}$$

It is then easy to see that:

$$\frac{\partial\widetilde{ca}_t}{\partial\varepsilon_t} = -(1-\lambda)\kappa\frac{\rho(1-\kappa)+h(\kappa-\lambda)-\rho\kappa h(1-\lambda)}{(1-\kappa)(1-\rho\kappa)} < 0 \qquad \frac{\partial\widetilde{ca}_t}{\partial\varepsilon_t\partial\lambda} = \frac{\kappa[\rho(1-\kappa)+h(\kappa-\lambda)+(h-2\rho\kappa h)(1-\lambda)]}{(1-\kappa)(1-\kappa\rho)} > 0$$

as long as λ, ρ, h, κ and $\in [0,1)$ and $\kappa \geq \lambda$, which will be satisfied under any plausible parametrization of this model.[10] Given these parameter restrictions, one can clearly see that the impact response of a net output log difference shock upon impact is negative and that greater liquidity constraints make the impact response of this shock less negative, thus

[9] See Enders, Müller and Scholl (2011) for more details on this approach.

[10] Since $\kappa = \exp(\gamma - \mu)$ and $\gamma - \mu$ is unlikely to be large, Kano (2008) argues that κ should be fairly close to, but smaller than, one. We therefore assume that $\kappa \in [0.9, 1)$ Since λ, the fraction of the population that is liquidity constrained, is unlikely to exceed this lower bound in reality, the condition $\kappa \geq \lambda$ will always be satisfied for any plausible parametrisation of the model.

smaller. In other words, financial deregulation, ie. the removal of liquidity constraints, makes the response of the current account to a net output log difference shock larger. Since the coefficient on $\widetilde{ca}_{t-1}$ is a function of λ one can also clearly see that the effect of a past shock on the current account today declines with a greater fraction of liquidity constrained agents. This means that the persistence of the current account is also decreasing in the liquidity constraint.

Under the assumption of internal habits and considering a net output log difference shock, the current account reaction function is

$$\begin{aligned} \widetilde{ca}_t &= h\widetilde{ca}_{t-1} + \lambda\widetilde{fb}_t - \lambda h\widetilde{fb}_{t-1} + \frac{(1-\lambda)\kappa(h-\rho)}{1-\rho\kappa}\Delta\widetilde{\ln NO}_{t-1} \\ &\quad -(1-\lambda)\kappa\frac{\rho(1-\kappa)+h\kappa(1-\rho)}{(1-\kappa)(1-\rho\kappa)}\varepsilon_t + f_t \end{aligned} \tag{13}$$

With regard to the size prediction, it is then easy to see that the effects of financial regulation under internal habits are qualitatively identical to those under external habits:

$$\frac{\partial\widetilde{ca}_t}{\partial\varepsilon_t} = -(1-\lambda)\kappa\frac{\rho(1-\kappa)+h\kappa(1-\rho)}{(1-\kappa)(1-\rho\kappa)} < 0 \qquad \frac{\partial\widetilde{ca}_t}{\partial\varepsilon_t\partial\lambda} = \kappa\frac{\rho(1-\kappa)+h\kappa(1-\rho)}{(1-\kappa)(1-\rho\kappa)} > 0$$

But unlike in the case of external habits, liquidity constraints do not affect current account persistence anymore. Let us turn to the case of net output shocks in log levels. In this case, the model with external habits becomes

$$\begin{aligned} \widetilde{ca}_t &= (1-\lambda)h\widetilde{ca}_{t-1} + \lambda\widetilde{fb}_t - \frac{(1-\lambda)\kappa(h-\rho)}{1-\rho\kappa}(1-\rho)\widetilde{\ln NO}_{t-1} \\ &\quad +(1-\lambda)\frac{((1-\kappa)+h\kappa(1-\rho)(1-\lambda))}{(1-\kappa)}\varepsilon_t + f_t \end{aligned} \tag{14}$$

It is then easy to show that $\frac{\partial\widetilde{ca}_t}{\partial\varepsilon_t} > 0$ and $\frac{\partial\widetilde{ca}_t}{\partial\varepsilon_t\partial\lambda} < 0$.

For the case of shocks in log-levels, the current account under internal habits is

$$\begin{aligned} \widetilde{ca}_t &= h\widetilde{ca}_{t-1} + \lambda\widetilde{fb}_t - \lambda h\widetilde{fb}_{t-1} - (1-\rho)\frac{(1-\lambda)\kappa(h-\rho)}{1-\rho\kappa}\widetilde{\ln NO}_{t-1} \\ &\quad +(1-\lambda)\frac{(1-\rho)h\kappa+(1-\kappa)}{(1-\kappa)}\varepsilon_t + f_t \end{aligned} \tag{15}$$

and

$$\frac{\partial \widetilde{ca}_t}{\partial \varepsilon_t} > 0 \quad \frac{\partial \widetilde{ca}_t}{\partial \varepsilon_t \partial \lambda} < 0$$

This means that the current account goes into surplus following a log level net output shock and that the reaction of the current account is smaller as liquidity constraints increase. As in the case of log difference shocks, with external habit formation, the persistence of the current account is declining in the liquidity constraints. In short, the model predicts that under lower liqudity constraints, the response of the current account becomes larger in any case, and more persistent in the presence of external habits, for a given log level/difference net output shock.

Finally note that our theory suggests that the data generating process for the current account will be a reduced form VAR model with $\widetilde{ca}_t$, $\widetilde{fb}_t$, $\Delta \widetilde{\ln NO}_t$ as the regressors.

3. Empirical methodology and data

This section develops a methodology to test the main implications of our model, namely that the reaction of the current account to net output shocks is larger and more persistent under lighter financial regulation.

3.1. General methodology

In practice, all the variables required to test this model are likely to be endogenous with respect to each other. To adress this concerns, we adopt a panel VAR approach where we explicitly identify net output shocks while addressing endogeneity concerns.[11] We use annual data given that financial regulation indices are not available at quarterly frequency. The panel VAR includes the three variables that drive the data-generating process for the current account according to our model: the current account to net output ratio, the fiscal balance to net output ratio and the net output growth rate. We therefore estimate the following interacted panel VAR model:

[11] An alternative approach would of course be the instrumental variable approach, but for annual data Chinn and Prasad (2003) have shown that it is difficult to obtain reliable instruments for the variables driving the current account.

$$A_0 Y_{i,t} = \gamma_t + \sum_{k=1}^{L} A_k Y_{i,t-k} + \sum_{k=0}^{L} B_k Y_{i,t-k} X_{i,t-k} + C X_{i,t} + e_{i,t} \quad e_{i,t} \sim N(0, \Sigma) \tag{16}$$

where $Y_{i,t}$ is an $n \times 1$ vector consisting of the current account to net output ratio, the fiscal balance to net output ratio and the net output growth rate in country i at time t. Since the theory provides predictions for variables expressed in deviations from their steady state value, we express these three variables in deviations from their country-specific means (which are our proxies for their steady state values). This removes the country fixed effect for these variables. γ_t is a time fixed effect.[12] A_k is an autoregressive matrix up to lag L. $e_{i,t}$ is an $n \times 1$ vector of residuals, assumed to be uncorrelated across countries and normally distributed with a constant constant covariance matrix Σ. The model is estimated with 2 lags.[13]

$X_{i,t}$ is a vector of the variables that we interact the panel VAR coefficients with. In other words, we allow the impact and lagged VAR coefficients to vary with the interaction term. Most existing studies model time variation in the VAR coefficients as a random walk (Cogley and Sargent, 2005 and Primiceri, 2005). In contrast to these papers, we follow Towbin and Weber (2011) and Sa, Towbin and Wieladek (2011) and allow the coefficients to vary with observable deterministic variables.[14] In our baseline specification, the $q \times 1$ vector $X_{i,t}$ contains $Finreg_{i,t}$, an index of financial regulation, and $KAOpen_{i,t}$, a measure of capital account openness. Previous work by Lewis (1997) has documented that restrictions on international transactions may lead consumers to act as if they are liquidity constrained. Similarly, Towbin (2008) has shown that the trade balance, an important component of the current account, becomes more persistent with greater capital account openess. To avoid omitted variable bias, capital account openess is therefore included as an independent determinant of the VAR coefficients. We use the *dejure* versions of these indicators because they are unlikely to be affected by net output shocks directly and can therefore be treated

[12]Since we want to ensure that we identify net output shocks of domestic, rather than external, origin, we include time fixed effects in the model to account for them.

[13]Ex-ante lag length selection criteria, such as the Akaike, Hanan-Quinn, and Schwartz-Bayesian criteria suggest a lag length of one. However, one the main assumptions of the VAR model is that residuals behave like white noise. Estimated with one lag, the residuals were autocorrelated of order 1, which is obviously inconsistent with white noise behaviour and suggested that in at least one of the equations 2 lags would be necessary. Since the bias from omitting a lag is typically worse than that from including an extra lag, we estimate the model with 2 lags.

[14]Our exposition of this model closely follows Sa, Towbin and Wieladek (2011).

as exogenous with respect to the VAR system.[15] In addition, $X_{i,t}$ enters separately to ensure that the change in the VAR coefficients does not occur due to omission of the interaction variable in levels.

The matrix A_0 is an $n \times n$ matrix with ones on the diagonal. Any element above the diagonal is 0. As a result of the lower triangular form, the residuals, $e_{i,t}$, will be orthogonal to each other and the covariance matrix Σ will be diagonal. The lower triangular elements of this matrix are interacted with $X_{i,t}$ and take the following form: The coefficient in row j and column h is modelled as $A_{0,i,t}(j,h) = A_0(j,h)Y_{i,t}(h) + B_0^1(j,h)Y_{i,t}(h)Finreg_{i,t} + B_1^1(j,h)Y_{i,t}(h)KAOpen_{i,t}$, where $j < h$. $B_0^1(j,h)$ and $B_1^1(j,h)$ are both scalar coefficients that denote the marginal change in financial regulation and capital account openess on $A_0(j,h)$, respectively. By setting $Finreg_{i,t}$ at a high (low) value, one can therefore obtain $A_{0,i,t}(j,h)$, which would correspond to the impact coefficient in a country with high (low) financial regulation. Estimating the VAR equation by equation in this recursive form provides a simple way of allowing the reduced form covariance matrix of the VAR, $\widetilde{\Sigma}_{i,t} = A_{0,i,t}^{-1}\Sigma A_{0,i,t}^{-1\prime}$, to vary with financial regulation and capital account openess. All of the coefficients in the lagged dependent variable coefficient matrix, $A_{k,i,t}(j,h)$, are also allowed to vary with the interaction terms, $X_{i,t}$, in an identical manner, but without any zero or one restrictions.

In summary, we are allowing the coefficients in both the impact and lagged dependent variable matrix to vary with observable deterministic variables. The advantage of this approach is that we can assess to which extent the impulse responses to a given shock differ with the degree of financial regulation (the liquidity constraint). Since our theory predicts that the effect of a net output shock on the current account to net output ratio depends on the degree of the liqudity constraint, this method is better suited to test our theory than standard time-varying coefficient VARs, which typically do not provide information on the source of time variation.

3.2. Identification of net output shocks

Previous work by Kano (2008) used zero restrictions derived from a model similar to (10) to identify net output shocks in Structural VARs. It is however difficult to establish whether the zero restrictions proposed by either of the models presented here are valid in the data. We therefore adopt an identification procedure that does not rely on zero restrictions. As we have shown in section 2.1, a log level net output shock increases the log level of net output

[15] These variables are not demeaned as they are already standardized.

and results in a current account surplus. Conversely, a log difference net output shock results in an increase in the log level of net output and a current account deficit. This provides mutually exclusive sign restrictions, robust across all of the theoretical models considered in this paper. We therefore use sign restrictions as pioneered by Canova and De Nicolo (2002), Faust and Rogers (2003) and Uhlig (2005) to identify the shocks of interest. In particular, we use the QR decomposition approach presented in Rubio-Ramirez, Waggoner, and Zha (2010) to search across the space of all possible structural VAR decompositions. Those that do not produce impulse responses that satisfy the restrictions in table 1 are discarded, while all of the remaining ones are kept for inference.

Table 1

	Log Level Net Output shock	Log Difference Net Output shock
LN NO_t	≥ 0	≥ 0
$\frac{CA_t}{NO_t}$	≥ 0	≤ 0
$\frac{FB_t}{NO_t}$	?	?

The avid reader will note that our model is estimated on the log difference, but that we impose identification restrictions on the log level, of net output. This is necessary to disentangle the two shocks, as the log level net output shock only implies a positive sign upon impact in net output log difference space and we impose restrictions upon impact and 4 years thereafter.[16] In other words, in the case of net output, we are imposing sign restrictions on the cumulative impulse responses of the log difference of this variable. Paustian (2007) points out that this methodology will only recover the structural shocks if a sufficient number of restrictions is imposed. To further ensure that this is the case, we also impose 'shape' restrictions on our impulse responses to rule out shocks to net output that do not correspond with the theory.[17] In particular, a stationary autoregressive log level (difference) net output process has the property that the response to an unexpected one-time shock does not increase (decrease) over time. Similarly, the theory implies that the current account response to any shock does not increase over time in absolute value. We therefore also discard structural decompositions which impy impulse responses that do not satisfy these shape restrictions.

Our proposed sign restrictions also emerge robustly from more complex open economy DSGE models. Fournier and Koske (2010) use the 2 country New Open Economy Macroeconomics model of Ferrero, Gertler and Svensson (2010), albeit without sticky prices, to investigate the effect of temporary and permanent productivity shocks on the current

[16]We note that imposing sign restrictions upon impact only does not produce substantially different results.

[17]Identifying shocks with only sign, but no shape, restrictions yields similar results.

account, through the savings channel. They find that the only scenario in which a temporary (permanent) productivity shock leads to a current account deficit (surplus) is by setting by their intertemporal elasticity of substitution parameter to 10, a high and unrealistic value compared to previous work in the literature.[18] Otherwise, their model predicts a non-negative (non-positive) current account response following a temporary (permanent) productivity shock. Similarly, Enders and Müller (2009) study the impact of permanent technology shocks on net exports[19], allowing for both the savings and investment channel, in a variant of the Backus, Kehoe and Kydland (1994) international RBC model. Under the assumption of incomplete markets, which is strongly supported by their time series evidence, they find that a permanent productivity shock leads to a current account deficit. Finally, Enders, Müller and Scholl (2011) derive theoretically robust sign restrictions from a very general 2 country DSGE model. They assume a near-unit root process for their technology shock, meaning that it can be interpreted as permanent, and find that with a low trade price elasticity, a feature of consistent with the data (See Enders and Müller, 2009), a very large fraction of their net exports impulse responses display a negative reaction to their technology shock. This implies, that even in their very general 2-country DSGE model with very wide parameter intervals, the fact that a permanent techonology shock has a negative effect on the current account is theoretically robust to most parametrizations.[20] Clearly, a temporary (permanent) technology shock would raise the level (growth rate) of output. We therefore argue that the log level (difference) net output shock identified in our model probably reflects a temporary (permanent) productivity shock in these more complex DSGE models. If this is truly the case, then our proposed identification restrictions are consistent both with the standard intertemporal model of the current account as well as more general open economy DSGE models.

[18] In their survey of this parameter value in calibrated open economy DSGE models, Enders, Müller and Scholl (2011) report that the intertemporal elasticity of substitution is typically set between .5 and 1. Empirically, Guvenen (2006) and Gruber (2006) estimate the elasticity of substitution to be 1 and 2, respectively. On the other hand, Hall (1988) and Yogo (2004) conclude that tit is not statistically different from zero. While the empirical literature had therefore not reached a conclusive answer, an elasticity of 10 is a order of magnitude larger than what is supported by previous empirical and theoretical work.

[19] Net exports, or the trade balance, is linked to the current account through the following identify: $CA_t = TB_t + rB_t$.

[20] Calibrated for annual data and with a perfect unit-root in the process for the productivity shock, it is probable that all of their net exports impulse responses would react negatively to the permanent productivity shock.

3.3. Estimation and inference

There are different ways of estimating (16). One option is to pool the coefficients across countries, assuming identical autoregressive dynamics across all units. If that assumption is violated, the resulting dynamic heterogeneity bias will typically lead to an upward bias in the VAR coefficients (Nickell, 1981; Canova, 2007), meaning that it is easy to mistake a temporary effect of a shock for a permanent one. Alternatively, Pesaran and Smith (1995) propose the mean group estimator as a solution to the problem of heterogeneity in the lagged slope coefficients.[21] This model is implemented by estimating the VAR model country by country and then averaging the country-specific VAR estimates to obtain the panel estimate. Rebucci (2003) points out that in small time series, as in our case, mean group panel VAR estimates may be subject to serious small sample bias. We therefore follow his prescription and pool the autoregressive coefficients across countries. However, since we allow the VAR coefficients to depend on financial regulation, the dynamic heterogeneity bias should not be a significant problem in our case: Our theoretical model predicts that the variation in the persistence of the current account is a result of different liquidity constaints across countries and our empirical strategy controls for this explicitly.

Following Uhlig (2005), we compute Bayesian error bands using a Normal-Wishart prior.[22] Since the covariance matrix in model (16) is diagonal by construction, we can proceed equation by equation. For each equation, we draw the parameters jointly from the posterior. Given the posterior draw, we then evaluate the interaction terms at the value of interest. For example, to obtain coefficients of the impact matrix, A_0, for a country with high financial regulation, we use $A_{0,FinregHigh}(j,h) = A_0(j,h) + B_0^1(j,h)Finreg^{HIGH} + B_1^1(j,h)KAOpen^{MEDIAN}$, where $Finreg^{HIGH}$ and $KAOpen^{MEDIAN}$ are high and median values of these indices. We use an analagous expression for each coefficient in the lagged dependent variable matrices. We then implemenent our identification procedure in the following manner: Given $A_{0,FinregHigh}$, which would correspond to the inverse of the Choleski decomposition in a reduced from VAR, we follow Rubio-Ramirez, Waggoner, and Zha (2010) and obtain the QR decomposition of a matrix of random numbers. We then verify if the structural decomposition $A_{0,FinregHigh}^{-1}Q$, together with the correspondingly evaluated lagged dependent variable matrices, generates impulse responses which are consistent with our sign and shape restirctions. If this is the case, we keep this draw of $A_{0,FinregHigh}^{-1}Q$ and the cor-

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[21] See Sa, Towbin and Wieladek (2011) for an application of the mean group estimator in interacted panel VAR models.

[22] Specifically, we implement algorithm 2.1 described in Del Negro and Schorfheide (2010). As in Cogely and Sargent (2005), we discard draws that imply explosive VAR dynamics.

responding impulse responses. We repeat this procedure until we have retained 100 draws which are used for inference.

3.4. Data

Our empirical strategy requires standard macroeconomic data and a good empirical measure of financial regulation. There are a few data sets that measure financial liberalization across countries (Williamson and Mahar (1998), Kaminsky and Schmukler (2003), Abiad et al (2010)). We rely on the data set by Abiad et al (2010) for basically two reasons. First, their country and time coverage is very wide (91 countries over the period 1973-2005) and thus appropriate for our panel VAR methodology. Second, their data set has seven graded components with special emphasis on domestic financial reform. Previous indices (eg, Kaminsky and Schmukler (2003)) put more weight on the liberalization of capital flows, which is not the central object of study in this paper. The seven components of the dataset are: credit controls, interest rate controls, entry barriers, state ownership in the banking sector, prudential regulation, securities market policy and capital account restrictions. Each component can take the values {0,1,2,3} with higher values meaning less regulation/restrictions.[23] We sum all components, but capital account restrictions, to come up with the aggregate domestic financial regulation index we use in our empirical exercise. This index is normalised to 1. Capital account restrictions are accounted for separately and we use the index by Chinn and Ito (2007).[24]

Data on the following macroeconomic variables for the period 1973-2005 is necessary to test the implications of the model: the current account balance, the fiscal balance, and a measure of net output. The IMF WEO gives us the current account balance and the fiscal balance in current US dollars. Net output is defined as GDP minus government consumption and investment. We obtain the shares of goverment consumption and investment in GDP from the World Bank WDI and we combine them with GDP in current US dollars from the IMF WEO to calculate net output. To obtain real net output per capita, we divide this by the US GDP deflator and total population. We remove outliers from the log difference of per capita net output, current account to net output ratio and fiscal balance to net output

[23] In the work by Abiad et al (2010), the prudential regulation component is an exception, as higher values reflect more prudential rules. Since we want higher values of the aggregate index to reflect greater financial regulation, we change the numbering in this category such that higher values indicate less regulation. We note that our results are robust to this change.

[24] Using the capital account openess index provided by Abiad et al (2010) yields similar results.

ratio by dropping values outside the 98 and 2 percentile of each variables distribution. In total, our sample consists of 84 countries (table 2) and 1756 annual observations.

4. Empirical results

In this section we present the results from our panel VAR model and explore a number of robustness checks. Evaluating the $Finreg_{i,t}$ term at high and low values of financial regulation permits us to obtain average VAR coefficients under regimes of high and low financial regulation. The remaning interaction term, $KAOpen_{i,t}$, is evaluated at the median value of its distribution. We can then obtain impulse responses under both regimes and compare them to assess whether the reaction of the current account varies with the financial regulation regime as predicted by the theory. The impulse response for the log level of net output was normalized to 1 in order to ensure that we are comparing current account responses to log level net output shocks of identical size across both regimes. This allows us to asses whether there is a statistically significant difference between the impulse responses due to a change in the financial regulatory regime, rather than the nature of net output shocks. Figure 4 shows impulse responses to a log level net output shock with the financial regulation term evaluated at both the 90% percentile ('low finreg' column) and 10% percentile ('high fin reg' column) of the distribution of this variable. These percentiles correspond to .23 and .83, respectively. The third column shows impulse responses obtained from a distribution of the difference in the impulse responses obtained under the high and low financial regulatory regime. The red and green dashed lines indicate the 90% and 68% confidence bands, respectively. Figure 5 repeats this exercise but for a log difference net output shock.

The median impact response of a log level net output shock on the current account is 0.57 under high financial regulation and 0.9 under low financial regulation, which represents a 58% increase in the impact of the shock if a country switches from high to low financial regulation (figure 4). As one can see in column three of figure 4, this difference is statistically significant. Furthermore looking at column three in figure 4, one can see that the median of the log net output response is close to 0 throughout. This suggests that the difference in the current account to net output response can not be attributed to changes in the nature of the net output shock. For a log difference net output shock (figure 5), the median impact response is -0.65 under high regulation and -1.05 under low regulation (61% increase). Again, the third column suggests that the difference in impulse responses is statistically significant and that the net output impulse responses are statistically not very different from each other. In both cases, the absolute value of the current account response is larger and more persistent under low financial regulation. Our theory predicted, that all else equal, the change in financial regulation should affect the impulse response of the current account balance to either type of shock in a similar way. Indeed, the change in the impact response and the persistence profile is very similar across both shocks, which provides additional verification for the theory.

We also evaluate the capital account openness term ($KAOpen$) at the 90% percentile ('high KA Open' column) and 10% percentile ('low KA Open' column) and the financial regulation term at the median. This yields less clear-cut results than for the financial regulation term. As figure 6 shows, the response of the current account to a log level net output shock intensifies at high levels of capital account restrictions but the third column shows that the difference between low and high capital account restrictions is not statistically significant. For the case of a log difference net output shock (figure 7), we find a more statistically significant role for capital account restrictions. But the difference in impulse responses goes in the wrong direction and as shown below is not robust.

4.1. Robustness

This subsection shows that the results presented so far are robust to the inclusion of the exchange rate regime as an additional determinant of the panel VAR coefficients and also attempts to adress potential concerns regarding dynamic heterogeneity bias.

The empirical results presented so far rely on the assumption that financial regulation and capital account restrictions are the only relevant determinants of the VAR coefficients. But there could also be other important empirical determinants of the VAR cofficients that we are not controlling for, leaving the findings presented thus far vulnerable to omitted variable bias. Another natural determinant of current account dynamics is the exchange rate regime. In theory, one would expect that the speed of adjustment of the current account under a fixed and floating exchange rate regime would differ (Friedman, 1953). However, Chinn and Wei (2011) document that it is difficult to find a statistically significant relationship between current account persistence and the exchange rate regime. Nevertheless, we also introduce the fine *defacto* exchange rate classification by Ilzetzki, Reinhart and Rogoff (2010) as an additional determinant of the VAR coefficients.[25] Figures 8-11 show that our previous results regarding the effect of financial regulation on the current account to net output ratio are robust to controlling for the exchange rate regime. Furthermore, the difference in the current account response between high and low capital account restrictions, to either type of shock, is not statistically significant anymore.

[25] Results using the de jure classification from the IMF's AREAER database are very similar.

The potential for dynamic heterogeneity bias in our panel VAR makes inference about the adjustment dynamics of the current account, with absolute certainty, difficult. Recently, Jarocinski (2010) introduced a new Bayesian panel VAR estimator to deal with this problem. In his approach, there is a shrinkage parameter λ which determines the degree of pooling in the panel VAR model. If $\lambda = 0$ then the VAR estimates will correspond to a pooled panel VAR model as in our case and if $\lambda \to \infty$, his model will estimate country-by-country VARs. The main advantage of his approach is that λ, and therefore the degree to which pooling is supported or not, is determined by the data. The disadvantage of his methodology is that it is not possible to estimate panel VARs with different variables in each equation, as (16). Nevertheless, to assess if dynamic heterogeneity bias is likely to be a problem in our application, we estimated the reduced form version of (16) with the Jarocinski (2010) approach, dropping all of the contemporaneous endogenous variables and the corresponding interaction terms from the model. We estimated this model with 150,000 repetitions of the Gibbs sampling chain, discarding the first 100,000 draws as burn in and keeping every 100th draw of the remainder for inference. While it is not possible to undertake similar structural VAR exercises with estimates from this model, the estimated λ is still likely to be informative to which extent pooling is appropriate. Jarocinski (2010) suggest looking at the square root of λ for this purpose. Figure 12 shows the distribution of $\sqrt[2]{\lambda}$ from this model. $\sqrt[2]{\lambda}$ is heavily clustered around a value of .0014, which is close to zero. This suggests that dynamic heterogeneity bias is probably not a substantial problem in our application.

Finally, there is also an issue regarding the potential observational equivalence of the net output shock in log differences and a stationary shock to the fiscal balance, $\widetilde{fb}_t$. Following a stationary negative shock to the fiscal balance, our theory predicts a current account deficit. If this shock has a positive effect on the log level of net output as well, it would not be possible to separately identify it from a log difference net output shock using only sign restrictions. But, since the governments budget has to be balanced in the long run, a temporary fiscal balance shock should only have a temporary effect on the log level of net output, with the maximum reached at impact and then declining over time. On the other hand, our shape restrictions require that log net output does not decrease over time, following a log difference net output shock. This allows us to differentiate these two shocks explicitly. Furthermore, our theory suggest that, all else equal, the easing of the liquidity constraint should have a similar effect on the current account, regardless of the assumed process for the log net output shock. Indeed, the results do suggest that this is the case, which implies that our proposed shape restriction is probably ensuring that inference from the retained impulse responses is not contaminated by observationally equivalent fiscal policy shocks.

5. Conclusion

Global current account imbalances and financial regulation have been at the forefront of economic policy debates since the global financial crisis broke out in late 2008. A growing literature has analyzed the role of global imbalances in the run-up to the crisis and the debate on the future regulatory landscape is equally active. However, the idea that financial regulation might affect the adjustment of current account imbalances remains largely unexplored despite its potentially important implications. This paper provides the first empirical evidence supporting the existence of a link between the degree of financial regulation and current account adjustment.

We introduce liquidity constraints, as a proxy for the degree of domestic financial regulation, into the standard intertemporal model of the current account and show that the response of the current account to a net output shock increases as fewer households are liquidity constrained. This conclusion is robust to various standard modeling choices, such as the introduction of a stochastic world real interest rate or habit formation in consumption. In the case of external habit formation in consumption, the persistence of the current account response increases with smaller liquidity constraints as well. Using a sample of 84 countries over the period 1973-2005, we test these theoretical predictions in an interacted panel VAR framework, identifying net output shocks with sign and shape restrictions. The structural VAR coefficents are allowed to vary with the degree of domestic financial regulation at the individual-country level. This feature allows us to assess if the degree of financial regulation affects the size and persistence response of the current account to a net output shock as predicted by the theory. In our baseline specification, the median current account impulse response to a net output shock is 60% larger upon impact, as well as substantially more persistent, under a low than under a high financial regulatory regime. This is robust to allowing the coefficients to vary with other potential determinants of current account adjusment, such as capital account openess and the exchange rate regime.

Exchange rate policy is typically blamed for the emergence of persistent and large current account imbalances. We find that the degree of domestic financial regulation has an independent, and quantitatively important, effect on the size and dynamics of current account balances. To our knowledge, this insight is not a prominent part of the current global financial reform debate but could have important implications for the future adjustment of global current account imbalances.

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A. Appendix

A.1. Deriving the linearized budget constraint

First, to go from (6) to (7), start with (6) :

$$E_t \sum_{i=0}^{\infty} R_{t,i} C_{t+i}^R = (1+r_t)(B_t^P + B_t^G) + E_t \sum_{i=0}^{\infty} R_{t,i} NO_{t+i}$$

then note that $R_{t,1} = 1$, factor out C_t^R and NO_t on the LHS and RHS, respectively:

$$C_t^R [1 + \sum_{i=1}^{\infty} E_t R_{t,i} \frac{C_{t+i}^R}{C_t^R}] = (1+r_t)(B_t^P + B_t^G) + NO_t [1 + \sum_{i=1}^{\infty} E_t R_{t,i} \frac{NO_{t+i}}{NO_t}]$$

Then divide the whole expression by NO_t

$$\frac{C_t^R}{NO_t} [1 + \sum_{i=1}^{\infty} E_t R_{t,i} \frac{C_{t+i}^R}{C_t^R}] = (1+r_t)(\frac{B_t^P + B_t^G}{NO_t}) + [1 + \sum_{i=1}^{\infty} E_t R_{t,i} \frac{NO_{t+i}}{NO_t}]$$

Now recall that $1/(\prod_{j=t+1}^{t+i} (1+r_j))$ and take logs and exponential in the infinite summations:

$$\begin{aligned} & \frac{C_t^R}{NO_t} [1 + \sum_{i=1}^{\infty} E_t \exp \left\{ \ln C_{t+i}^R - \ln C_t^R - \ln(\prod_{j=t+1}^{t+i} (1+r_j)) \right\}] \\ = \quad & \exp(\ln(1+r_t))(\frac{B_t^P + B_t^G}{NO_t}) + [1 + \sum_{i=1}^{\infty} E_t \exp \left\{ \ln NO_{t+i} - \ln NO_t - \ln(\prod_{j=t+1}^{t+i} (1+r_j)) \right\}] \end{aligned}$$

Applying the log to the interest rate factorial, it is easy to see that:

$$\begin{aligned} & \frac{C_t^R}{NO_t} [1 + \sum_{i=1}^{\infty} E_t \exp \left\{ \ln C_{t+i}^R - \ln C_t^R - (\sum_{j=t+1}^{t+i} \ln(1+r_j)) \right\}] \\ = \quad & \exp(\ln(1+r_t))(\frac{B_t^P + B_t^G}{NO_t}) + [1 + \sum_{i=1}^{\infty} E_t \exp \left\{ \ln NO_{t+i} - \ln NO_t - (\sum_{j=t+1}^{t+i} \ln(1+r_j)) \right\}] \end{aligned}$$

Now we can add and subtract infinitely many $\ln C_t^{R\prime} s$ and $\ln NO_t$'s:

$$
\begin{aligned}
& \frac{C_t^R}{NO_t}[1+\sum_{i=1}^{\infty} E_t \exp \left\{ \begin{array}{c} \ln C_{t+i}^R - \ln C_{t+i-1}^R + \ln C_{t+i-1}^R \cdots \\ -\ln C_{t+1}^R + \ln C_{t+1}^R - \ln C_t^R - (\sum_{j=t+1}^{t+i} \ln(1+r_j)) \end{array} \right\}] \\
= \ & \exp(\ln(1+r_t))(\frac{B_t^P + B_t^G}{NO_t}) + \\
& [1+\sum_{i=1}^{\infty} E_t \exp \left\{ \begin{array}{c} \ln NO_{t+i} - \ln NO_{t+i-1} + \ln NO_{t+i-1} \cdots \\ -\ln NO_{t+1} + \ln NO_{t+1} - \ln NO_t - (\sum_{j=t+1}^{t+i} \ln(1+r_j)) \end{array} \right\}]
\end{aligned}
$$

which is (7) in the main text:

$$
\begin{aligned}
& \frac{C_t}{NO_t}[1+\sum_{i=1}^{\infty} E_t \exp \left\{ \sum_{j=t+1}^{t+i} (\Delta \ln C_j - \ln(1+r_j)) \right\}] \\
= \ & \exp\{\ln(1+r_t)\} \frac{B_t}{NO_t} + [1+\sum_{i=1}^{\infty} E_t \exp \left\{ \sum_{j=t+1}^{t+i} (\Delta \ln NO_j - \ln(1+r_j)) \right\}]
\end{aligned}
$$

To linearize this expression, we use the standard formula $f(x) = f(a) + f'(x)(x-a)$

$$
\begin{aligned}
& c[1+\sum_{i=1}^{\infty}\kappa^i] + \widetilde{\frac{C_t}{NO_t}}[1+\sum_{i=1}^{\infty}\kappa^i] + c[1+\sum_{i=1}^{\infty}\kappa^i]\sum_{i=1}^{\infty}\kappa^i E_t \left\{ \Delta \widetilde{\ln C^R}_{t+i} - \ln \widetilde{(1+r_{t+i})} \right\} \\
= \ & \frac{1}{e^{-\mu}}\frac{B_t}{NO_t} + \frac{b}{e^{-\mu}}\widetilde{\ln(1+r_t)} + [1+\sum_{i=1}^{\infty}\kappa^i] + [1+\sum_{i=1}^{\infty}\kappa^i][\sum_{i=1}^{\infty}\kappa^i E_t \left\{ \Delta \widetilde{\ln NO}_{t+i} - \ln \widetilde{(1+r_{t+i})} \right\}]
\end{aligned}
$$

Now note that $[1+\sum_{i=1}^{\infty}\kappa^i] = [1+\kappa\sum_{i=0}^{\infty}\kappa^i] = [1+\frac{\kappa}{1-\kappa}] = [\frac{1}{1-\kappa}]$ and also note that in steady state, when all the hat variables take the value of 0, $c[1+\sum_{i=1}^{\infty}\kappa^i] = [1+\sum_{i=1}^{\infty}\kappa^i]$, then it is easy to obtain the final linearized form:

$$
\begin{aligned}
\widetilde{\frac{C_t^R}{NO_t}} \ = \ & \frac{1-\kappa}{e^{-\mu}}(\widetilde{\frac{B_t^G}{NO_t}} + \widetilde{\frac{B_t^P}{NO_t}}) - c\sum_{i=1}^{\infty}\kappa^i E_t \left\{ \Delta \widetilde{\ln C^R}_{t+i} - \ln \widetilde{(1+r_{t+i})} \right\} \\
& + \sum_{i=1}^{\infty}\kappa^i E_t \left\{ \Delta \widetilde{\ln NO}_{t+i} - \ln \widetilde{(1+r_{t+i})} \right\} + \frac{1-\kappa}{e^{-\mu}}(b^P + b^G)\ln \widetilde{(1+r_t)}
\end{aligned} \tag{A1}
$$

A.2. Derivation of the current account reaction function with external habits and a constant world real interest rate

We start by linearizing the definition of aggregate consumption $C_t = (1-\lambda)C_t^R + \lambda C_t^{NR}$ to obtain

$$\frac{\widetilde{C_t}}{NO_t} = (1-\lambda)\frac{\widetilde{C_t^R}}{NO_t} + \lambda(\frac{\widetilde{G_t}}{NO_t} - \frac{\widetilde{T_t}}{NO_t}). \tag{A2}$$

Following Obstfeld and Rogoff (1996), the current account can be expressed as as:

$$CA_t = B_{t+1} - B_t = NO_t + rB_t - C_t$$

which, noting that $B_t = (1-\lambda)B_t^P + B_t^G$ and $r = e^{\ln(1+r)} - 1$, can be linearized as the following the current account to net output ratio:

$$\widetilde{ca}_t = (e^{\mu}-1)((1-\lambda)\frac{\widetilde{B_t^P}}{NO_t} + \frac{\widetilde{B_t^G}}{NO_t}) - \frac{\widetilde{C_t}}{NO_t} \tag{A3}$$

where $\widetilde{ca}_t = \frac{\widetilde{CA_t}}{NO_t}$. Now plugging (A1) into (A2), then into (A3) and simplifying yields[26]:

$$\begin{aligned}
\widetilde{ca}_t \;\; = \;\; & \lambda[r\frac{\widetilde{B_t^G}}{NO_t} + \frac{\widetilde{T_t}}{NO_t} - \frac{\widetilde{G_t}}{NO_t}] - ((1-\lambda)(e^{\mu}-1)b^G - e^{\mu}b^G)\ln\widetilde{(1+r_t)} \\
& +(1-\lambda)\frac{\overline{C^R}}{\overline{NO}}\sum_{i=1}^{\infty}\kappa^i E_t\left\{\Delta\ln\widetilde{C^R}_{t+i} - \ln\widetilde{(1+r_{t+i})}\right\} \\
& -(1-\lambda)\sum_{i=1}^{\infty}\kappa^i E_t\left\{\Delta\ln\widetilde{NO}_{t+i} - \ln\widetilde{(1+r_{t+i})}\right\}
\end{aligned}$$

The constant world real interest rate assumption implies $\ln\widetilde{(1+r_t)} = 0 \; \forall \; i$. We also add and subtract $(1-\lambda)h\widetilde{ca}_{t-1}$ and define $\widetilde{fb}_t = [r\frac{\widetilde{B_t^G}}{NO_t} + \frac{\widetilde{T_t}}{NO_t} - \frac{\widetilde{G_t}}{NO_t}]$ to obtain

---

[26] Where we used $\frac{1}{e^{\gamma}} = 1$ and $(e^{\mu}-1) = r$ as in Kano (2008).

$$
\begin{aligned}
\widetilde{ca}_t &= (1-\lambda)h\widetilde{ca}_{t-1} + \lambda fb_t + (1-\lambda)\frac{\overline{C^R}}{\overline{NO}}\sum_{i=1}^{\infty}\kappa^i E_t\left\{\Delta\widetilde{\ln C^R}_{t+i}\right\} \qquad \text{(A4)}\\
&-(1-\lambda)h\lambda fb_{t-1} - h(1-\lambda)^2\frac{\overline{C^R}}{\overline{NO}}\sum_{i=1}^{\infty}\kappa^i E_{t-1}\left\{\Delta\widetilde{\ln C^R}_{t+i-1}\right\}\\
&-(1-\lambda)\sum_{i=1}^{\infty}\kappa^i E_t\left\{\Delta\widetilde{\ln NO}_{t+i}\right\} + h(1-\lambda)^2\sum_{i=1}^{\infty}\kappa^i E_{t-1}\left\{\Delta\widetilde{\ln NO}_{t+i-1}\right\}
\end{aligned}
$$

Furthermore, log-linearizing the Euler equation gives

$$
E_t\Delta\widetilde{\ln C^R}_t = h\frac{\overline{C}}{\overline{C^R}}\Delta\widetilde{\ln C}_{t-1}
$$

which can be expressed as

$$
E_t\Delta\widetilde{\ln C^R}_t = h(1-\lambda)\Delta\widetilde{\ln C^R}_{t-1} + h\lambda\frac{\overline{NO}}{\overline{C^R}}\Delta\widetilde{\ln NO}_{t-1} + h\lambda\frac{\overline{G}}{\overline{C^R}}\Delta\widetilde{\ln G}_{t-1} - h\lambda\frac{\overline{T}}{\overline{C^R}}\Delta\widetilde{\ln T}_{t-1} \qquad \text{(A5)}
$$

Plugging (A5) into (A4) gives

$$
\begin{aligned}
\widetilde{ca}_t &= (1-\lambda)h\widetilde{ca}_{t-1} + \lambda\widetilde{fb}_t + h(1-\lambda)^2 c\sum_{i=1}^{\infty}\kappa^i E_t\left\{\Delta\widetilde{\ln C^R}_{t+i-1}\right\} \qquad \text{(A6)}\\
&+h\lambda(1-\lambda)\sum_{i=1}^{\infty}\kappa^i E_t\left\{\Delta\widetilde{\ln NO}_{t+i-1}\right\} + h\lambda(1-\lambda)g\sum_{i=1}^{\infty}\kappa^i E_t\left\{\Delta\widetilde{\ln G}_{t+i-1}\right\}\\
&-h\lambda(1-\lambda)t\sum_{i=1}^{\infty}\kappa^i E_t\left\{\Delta\widetilde{\ln T}_{t+i-1}\right\} - (1-\lambda)h\lambda\widetilde{fb}_{t-1}\\
&-h(1-\lambda)^2 c\sum_{i=1}^{\infty}\kappa^i E_{t-1}\left\{\Delta\widetilde{\ln C^R}_{t+i-1}\right\} - (1-\lambda)\sum_{i=1}^{\infty}\kappa^i E_t\left\{\Delta\widetilde{\ln NO}_{t+i}\right\}\\
&+h(1-\lambda)^2\sum_{i=1}^{\infty}\kappa^i E_{t-1}\left\{\Delta\widetilde{\ln NO}_{t+i-1}\right\}
\end{aligned}
$$

where $c = \frac{\overline{C^R}}{\overline{NO}}$; $g = \frac{\overline{G}}{\overline{NO}}$ and $t = \frac{\overline{T}}{\overline{NO}}$.

To further solve the model, it is useful to derive the governments linearised budget constraint first. Using $E_t \sum_{i=0}^{\infty} R_{t,i} G_{t+i} = (1+r_t)(B_t^G) + E_t \sum_{i=0}^{\infty} R_{t,i} T_{t+i}$, the governments intertemporal budget constraint can be written as:

$$\frac{G_t}{NO_t}[1 + E_t \sum_{i=0}^{\infty} R_{t,i} \frac{G_{t+i}}{G_t}] = (1+r_t)(\frac{B_t^G}{NO_t}) + [1 + E_t \sum_{i=0}^{\infty} R_{t,i} T_{t+i}] \frac{T_t}{NO_t}$$

The linearized version of this expression is

$$\begin{aligned} & g[1+\sum_{i=1}^{\infty}\kappa^i] + \widetilde{\frac{G_t}{NO_t}}[1+\sum_{i=1}^{\infty}\kappa^i] + g[1+\sum_{i=1}^{\infty}\kappa^i]\sum_{i=1}^{\infty}\kappa^i E_t\left\{\Delta \widetilde{\ln G}_{t+i}\right\} \\ = \;\; & \frac{1}{e^{-\mu}}\widetilde{\frac{B_t^G}{NO_t}} + t[1+\sum_{i=1}^{\infty}\kappa^i] + \widetilde{\frac{T_t}{NO_t}}[1+\sum_{i=1}^{\infty}\kappa^i] + t[1+\sum_{i=1}^{\infty}\kappa^i][\sum_{i=1}^{\infty}\kappa^i E_t\left\{\Delta \widetilde{\ln T}_{t+i}\right\}] \end{aligned}$$

Under the assumption that $g = t$, this can then be expressed as :

$$\begin{aligned} & g\sum_{i=1}^{\infty}\kappa^i E_t\left\{\Delta \widetilde{\ln G}_{t+i}\right\} - t[\sum_{i=1}^{\infty}\kappa^i E_t\left\{\Delta \widetilde{\ln T}_{t+i}\right\}] \\ = \;\; & \frac{1-\kappa}{e^{-\mu}}\widetilde{\frac{B_t^G}{NO_t}} + \widetilde{\frac{T_t}{NO_t}} - \widetilde{\frac{G_t}{NO_t}} \end{aligned} \tag{A7}$$

Plugging (A7) into (A6) yields

$$\begin{aligned} \widetilde{ca}_t \;\; = \;\; & (1-\lambda)h\widetilde{ca}_{t-1} + \lambda\widetilde{fb}_t + h(1-\lambda)^2 c\sum_{i=1}^{\infty}\kappa^i(E_t - E_{t-1})\left\{\Delta \widetilde{\ln C^R}_{t+i-1}\right\} \\ & + h\lambda(1-\lambda)\sum_{i=1}^{\infty}\kappa^i E_t\left\{\Delta \widetilde{\ln NO}_{t+i-1}\right\} \\ & -(1-\lambda)\sum_{i=1}^{\infty}\kappa^i E_t\left\{\Delta \widetilde{\ln NO}_{t+i}\right\} + h(1-\lambda)^2\sum_{i=1}^{\infty}\kappa^i E_{t-1}\left\{\Delta \widetilde{\ln NO}_{t+i-1}\right\} \end{aligned}$$

Define $f_t = h(1-\lambda)^2 c\sum_{i=1}^{\infty}\kappa^i(E_t - E_{t-1})\left\{\Delta \widetilde{\ln C^R}_{t+i-1}\right\}$ and use $\sum_{i=1}^{\infty}\kappa^i E_t\left\{\Delta \widetilde{\ln NO}_{t+i-1}\right\} = \kappa\sum_{i=0}^{\infty}\kappa^i E_t\left\{\Delta \widetilde{\ln NO}_{t+i}\right\}$ and $\sum_{i=1}^{\infty}\kappa^i E_t\left\{\Delta \widetilde{\ln NO}_{t+i}\right\} + \Delta \widetilde{\ln NO}_t = \sum_{i=0}^{\infty}\kappa^i E_t\left\{\Delta \widetilde{\ln NO}_{t+i}\right\}$

to obtain

$$\begin{aligned}\widetilde{ca}_t &= (1-\lambda)h\widetilde{ca}_{t-1} + \lambda\widetilde{fb}_t + f_t - (1-h\kappa\lambda)(1-\lambda)\sum_{i=0}^{\infty}\kappa^i E_t\left\{\Delta\widetilde{\ln NO}_{t+i}\right\} \\ &+h(1-\lambda)^2\sum_{i=1}^{\infty}\kappa^i E_{t-1}\left\{\Delta\widetilde{\ln NO}_{t+i-1}\right\} + (1-\lambda)\Delta\widetilde{\ln NO}_t\end{aligned}$$

Now using $\sum_{i=1}^{\infty}\kappa^i E_t\left\{\Delta\widetilde{\ln NO}_{t+i-1}\right\} = \kappa\sum_{i=0}^{\infty}\kappa^i E_t\left\{\Delta\widetilde{\ln NO}_{t+i}\right\}$ and adding and subtracting $h\kappa(1-\lambda)^2\sum_{i=0}^{\infty}\kappa^i E_t\left\{\Delta\widetilde{\ln NO}_{t+i}\right\}$

$$\begin{aligned}\widetilde{ca}_t &= (1-\lambda)h\widetilde{ca}_{t-1} + \lambda\widetilde{fb}_t - (1-h\kappa)(1-\lambda)\sum_{i=0}^{\infty}\kappa^i E_t\left\{\Delta\widetilde{\ln NO}_{t+i}\right\} \\ &+(1-\lambda)\Delta\widetilde{\ln NO}_t - h\kappa(1-\lambda)^2\sum_{i=0}^{\infty}\kappa^i(E_t - E_{t-1})\left\{\Delta\widetilde{\ln NO}_{t+i}\right\} + f_t\end{aligned}$$

which is (9) in the main text. To further solve this expression, it is necessary to assume a process for $\widetilde{\ln NO}_t$. Under the assumption of $\Delta\ln NO_t = \rho\Delta\ln NO_{t-1} + \varepsilon_t$ (9) can be expressed as:

$$\begin{aligned}\widetilde{ca}_t &= (1-\lambda)h\widetilde{ca}_{t-1} + \lambda\widetilde{fb}_t - \frac{(1-\lambda)(1-h\kappa)}{1-\rho\kappa}\Delta\widetilde{\ln NO}_t \\ &+(1-\lambda)\Delta\widetilde{\ln NO}_t - \frac{h\kappa(1-\lambda)(1-\lambda)}{(1-\kappa)}\varepsilon_t + f_t\end{aligned}$$

where we replaced $h\kappa(1-\lambda)(1-\lambda)\sum_{i=0}^{\infty}\kappa^i(E_t - E_{t-1})\Delta\widetilde{\ln NO}_{t+i}$ with $\frac{h\kappa(1-\lambda)(1-\lambda)}{(1-\kappa)}\varepsilon_t$ and $(1-\lambda)(1-h\kappa)\sum_{i=0}^{\infty}\kappa^i E_t\Delta\widetilde{\ln NO}_{t+i}$ with $\frac{(1-\lambda)(1-h\kappa)}{1-\rho\kappa}\Delta\widetilde{\ln NO}_t$. After further simplification, one can then obtain (12) in the main text:

$$\begin{aligned}\widetilde{ca}_t &= (1-\lambda)h\widetilde{ca}_{t-1} + \lambda\widetilde{fb}_t - \frac{(1-\lambda)\kappa(\rho-h)}{1-\rho\kappa}\rho\Delta\widetilde{\ln NO}_{t-1} \\ &-\frac{\kappa(1-\lambda)(\rho(1-\kappa)+h(\kappa-\lambda)-\rho\kappa h(1-\lambda))}{(1-\kappa)(1-\rho\kappa)}\varepsilon_t + f_t\end{aligned}$$

To derive (14) note that under the assumption that $\ln NO_t = \rho\ln NO_{t-1} + \varepsilon_t$, $\Delta\widetilde{\ln NO}_t = (\rho-1)\ln NO_{t-1} + \varepsilon_t$ and follow the steps above. In this case $h\kappa(1-\lambda)(1-\lambda)\sum_{i=0}^{\infty}\kappa^i(E_t - E_{t-1})\Delta\widetilde{\ln NO}_{t+i} = -(1-\rho)\frac{h\kappa(1-\lambda)(1-\lambda)}{(1-\kappa)}\varepsilon_t$ and $(1-\lambda)(1-h\kappa)\sum_{i=0}^{\infty}\kappa^i E_t\Delta\widetilde{\ln NO}_{t+i} = -(1-\rho)\frac{(1-\lambda)(1-h\kappa)}{1-\rho\kappa}\widetilde{\ln NO}_{t-1}$.

A.3. Derivation of the current account reaction function with internal habits and a constant world real interest rate

Start with

$$\widetilde{ca}_t = \lambda \widetilde{fb}_t + (1-\lambda)\frac{c\kappa h}{1-\kappa h}\widetilde{\Delta \ln C^R}_t - (1-\lambda)\sum_{i=1}^{\infty}\kappa^i E_t\left\{\widetilde{\Delta \ln NO}_{t+i}\right\}$$

In order to solve out for the growth rate of consumption in the current account equation, we add and subtract $h\widetilde{ca}_{t-1}$ on both sides of the equation. This yields

$$\begin{aligned}
\widetilde{ca}_t &= h\widetilde{ca}_{t-1} + \lambda \widetilde{fb}_t - \lambda h \widetilde{fb}_{t-1} - (1-\lambda)\sum_{i=1}^{\infty}\kappa^i E_t\left\{\widetilde{\Delta \ln NO}_{t+i}\right\} \\
&+ (1-\lambda)h\sum_{i=1}^{\infty}\kappa^i E_{t-1}\left\{\widetilde{\Delta \ln NO}_{t+i}\right\} \\
&+ (1-\lambda)\frac{c\kappa h}{1-\kappa h}(\widetilde{\Delta \ln C^R}_t - h\widetilde{\Delta \ln C^R}_{t-1})
\end{aligned}$$

Now note that with internal habits in consumption, (4) can be log-linearized as $E_t\widetilde{\Delta \ln C^R}_{t+1} = h\widetilde{\Delta \ln C^R}_t$. Then use $\sum_{i=1}^{\infty}\kappa^i E_t\left\{\widetilde{\Delta \ln NO}_{t+i-1}\right\} = \kappa\sum_{i=0}^{\infty}\kappa^i E_t\left\{\widetilde{\Delta \ln NO}_{t+i}\right\}$, $\sum_{i=1}^{\infty}\kappa^i E_t\left\{\widetilde{\Delta \ln NO}_{t+i}\right\} + \widetilde{\Delta \ln NO}_t = \sum_{i=0}^{\infty}\kappa^i E_t\left\{\widetilde{\Delta \ln NO}_{t+i}\right\}$ and add and subtract $(1-\lambda)(h\kappa)\sum_{i=0}^{\infty}\kappa^i E_t\left\{\widetilde{\Delta \ln NO}_{t+i}\right\}$ to obtain (10) in the main text:

$$\begin{aligned}
\widetilde{ca}_t &= h\widetilde{ca}_{t-1} + \lambda \widetilde{fb}_t - \lambda h \widetilde{fb}_{t-1} + (1-\lambda)\widetilde{\Delta \ln NO}_t - (1-\lambda)(1-h\kappa)\sum_{i=0}^{\infty}\kappa^i E_t\left\{\widetilde{\Delta \ln NO}_{t+i}\right\} \\
&- (1-\lambda)(h\kappa)\sum_{i=0}^{\infty}\kappa^i (E_t - E_{t-1})\left\{\widetilde{\Delta \ln NO}_{t+i}\right\} + f_t
\end{aligned}$$

where $f_t = (1-\lambda)\frac{c\kappa h}{1-\kappa h}(\widetilde{\Delta \ln C^R}_t - E_{t-1}\widetilde{\Delta \ln C^R}_t)$. Solving under the assumption that $\Delta \ln NO_t = \rho \Delta \ln NO_{t-1} + \varepsilon_t$ yields (13)

$$\begin{aligned}
\widetilde{ca}_t &= h\widetilde{ca}_{t-1} + \lambda \widetilde{fb}_t - \lambda h \widetilde{fb}_{t-1} + \frac{(1-\lambda)\kappa(h-\rho)}{1-\rho\kappa}\widetilde{\Delta \ln NO}_{t-1} \\
&- (1-\lambda)\kappa\frac{\rho(1-\kappa) + h\kappa(1-\rho)}{(1-\kappa)(1-\rho\kappa)}\varepsilon_t + f_t
\end{aligned}$$

where we replaced $(1-\lambda)(1-h\kappa)\sum_{i=0}^{\infty}\kappa^i E_t\left\{\widetilde{\Delta\ln NO}_{t+i}\right\}$ with $\frac{(1-\lambda)(1-h\kappa)\rho\kappa}{1-\rho\kappa}\widetilde{\Delta\ln NO}_t$, using $\frac{1}{1-x}=\sum_{i=0}^{\infty}x^i$, and $(1-\lambda)(h\kappa)\sum_{i=0}^{\infty}\kappa^i(E_t-E_{t-1})\left\{\widetilde{\Delta\ln NO}_{t+i}\right\}$ with $\frac{(1-\lambda)(h\kappa)}{(1-\kappa)}\varepsilon_t$. To derive (15) note that under the assumption that $\ln NO_t = \rho\ln NO_{t-1}+\varepsilon_t$, $\widetilde{\Delta\ln NO}_t = (\rho-1)\ln NO_{t-1}+\varepsilon_t$ and follow the same steps as before.

A.4. Derivation of the current account reaction function under a stochastic time-varying world real interest rate and no habitual consumption

This section derives the current account reaction function under a stochastic world interest rate and shows the observational equivalence with the model with internal habits. As a first step, linearize the definition of the current account ($CA_t = NO_t + r_t B_t - C_t$) under the assumption of a time-varying stochastic world real interest rate to obtain

$$\widetilde{ca}_t = (e^{\mu}-1)((1-\lambda)\frac{\widetilde{B_t^P}}{NO_t}+\frac{\widetilde{B_t^G}}{NO_t}) + e^{\mu}((1-\lambda)b^P+b^G)\ln\widetilde{(1+r_t)} - \frac{\widetilde{C_t}}{NO_t} \tag{A8}$$

Equation (4) can be written as

$$E_t\Delta\ln C_{t+1}^R = \delta + \ln\beta + E_t\ln(1+r_{t+1})$$

linearizing this expression yields: $\widetilde{\Delta\ln C^R}_t = \ln\widetilde{(1+r_{t+i})}$. Plugging this into (8) yields

$$\frac{\widetilde{C_t^R}}{NO_t} = \frac{1-\kappa}{e^{-\mu}}(\frac{\widetilde{B_t^G}}{NO_t}+\frac{\widetilde{B_t^P}}{NO_t})+\frac{1-\kappa}{e^{-\mu}}(b^P+b^G)\ln\widetilde{(1+r_t)}+\sum_{i=1}^{\infty}\kappa^i E_t\left\{\widetilde{\Delta\ln NO}_{t+i}-\ln\widetilde{(1+r_{t+i})}\right\}$$

Now, plugging this into (A8) and simplifying[27] yields:

$$\begin{aligned}\widetilde{ca}_t &= \lambda[r\frac{\widetilde{B_t^G}}{NO_t}+\frac{\widetilde{T_t}}{NO_t}-\frac{\widetilde{G_t}}{NO_t}] - ((1-\lambda)(e^{\mu}-1)b^G - e^{\mu}b^G)\ln\widetilde{(1+r_t)} \\ &\quad -(1-\lambda)\sum_{i=1}^{\infty}\kappa^i E_t\left\{\widetilde{\Delta\ln NO}_{t+i}\right\} + (1-\lambda)\sum_{i=1}^{\infty}\kappa^i E_t\left\{\ln\widetilde{(1+r_{t+i})}\right\}\end{aligned}$$

since $e^{\gamma}\approx 1$.

Now, to show observational equivalence between the above and the model with internal habits, note that $\ln \widetilde{(1 + r_{t+i})} \equiv \widetilde{r}_{t+i}$. Assuming an AR(1) process for the world real interest rate, $\widetilde{r}_t = \rho^r \widetilde{r}_{t-1} + \eta_t$, one can then rewrite the above as

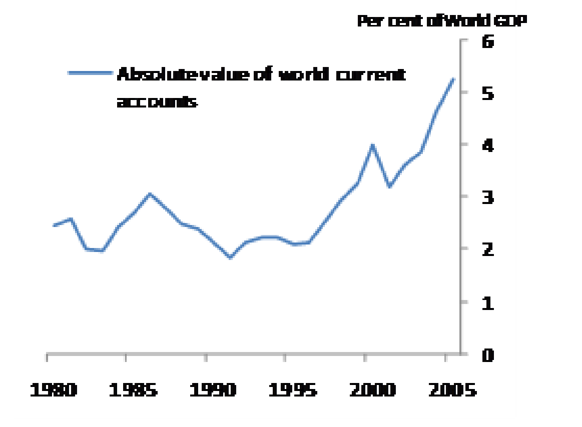
$$\begin{aligned} \widetilde{ca}_t &= \lambda[r \frac{\widetilde{B_t^G}}{NO_t} + \frac{\widetilde{T_t}}{NO_t} - \frac{\widetilde{G_t}}{NO_t}] - ((1-\lambda)(e^{\mu}-1)b^G - e^{\mu}b^G)\rho^r \widetilde{r}_{t-1} \\ &\quad -(1-\lambda)\sum_{i=1}^{\infty}\kappa^i E_t \left\{\Delta \widetilde{\ln NO}_{t+i}\right\} + \frac{(1-\lambda)\rho^r\kappa}{1-\rho^r\kappa}\rho^r \widetilde{r}_{t-1} + f_t \end{aligned} \tag{A9}$$

where $f_t = [((1-\lambda)(e^{\mu}-1)b^G - e^{\mu}b^G) + \frac{(1-\lambda)\rho^r\kappa}{1-\rho^r\kappa}]\eta_t$. Comparing these last two expressions and using $\sum_{i=1}^{\infty}\kappa^i E_t \left\{\Delta \widetilde{\ln NO}_{t+i}\right\} + \Delta \widetilde{\ln NO}_t = \sum_{i=0}^{\infty}\kappa^i E_t \left\{\Delta \widetilde{\ln NO}_{t+i}\right\}$, one can see that:

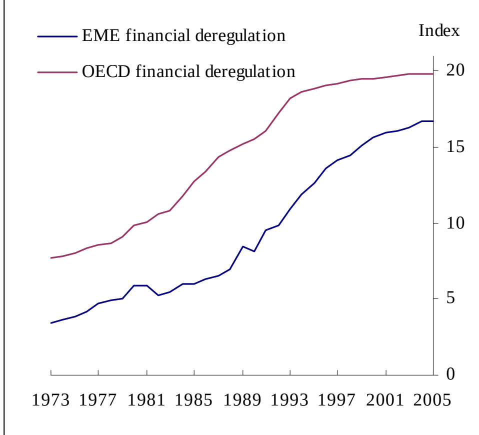
$$\begin{aligned} \widetilde{ca}_t &= \rho^r \widetilde{ca}_t + \lambda[r \frac{\widetilde{B_t^G}}{NO_t} + \frac{\widetilde{T_t}}{NO_t} - \frac{\widetilde{G_t}}{NO_t}] - \rho^r \lambda[r \frac{\widetilde{B_{t-1}^G}}{NO_{t-1}} + \frac{\widetilde{T_{t-1}}}{NO_{t-1}} - \frac{\widetilde{G_{t-1}}}{NO_{t-1}}] + (1-\lambda)\Delta \widetilde{\ln NO}_t \\ &\quad -(1-\lambda)(1-\rho^r\kappa)\sum_{i=0}^{\infty}\kappa^i E_t \left\{\Delta \widetilde{\ln NO}_{t+i}\right\} - (1-\lambda)(\rho^r\kappa)\sum_{i=0}^{\infty}\kappa^i (E_t - E_{t-1})\left\{\Delta \widetilde{\ln NO}_{t+i}\right\} + f_t \end{aligned}$$

which, as in Kano (2009), is observationally equivalent to the solution under internal habits.

Figure 1	Figure 2
Size of current account imbalances over time	Financial deregulation over time



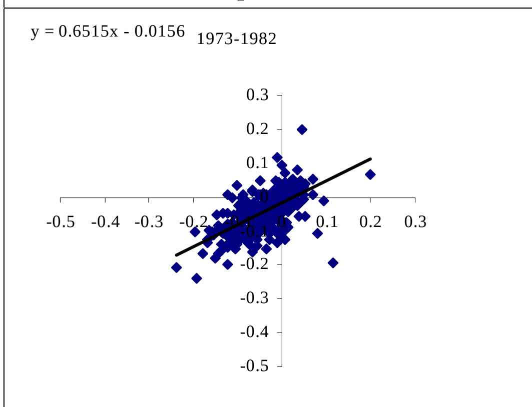
Source: IMF WEO.



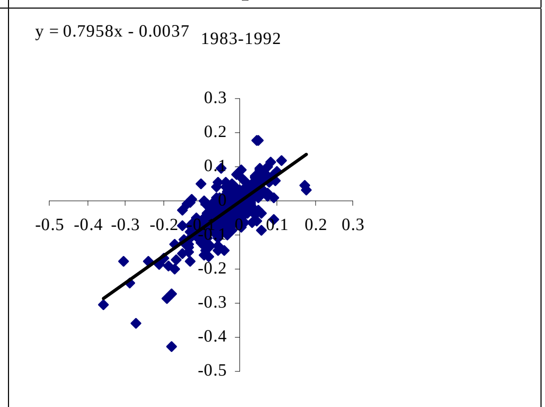
Source: Abiad et al (2010).

Figure 3

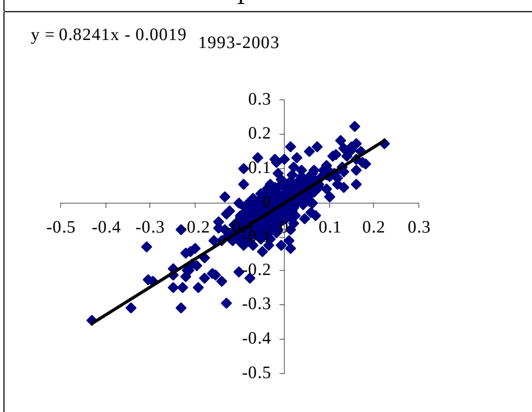
Current account persistence 1973 -1982



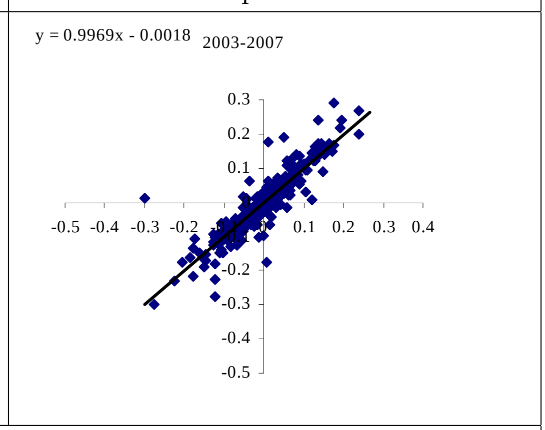
Current account persistence 1983 -1992



Current account persistence 1993 -2003



Current account persistence 2003 -2007



Source: Authors calculation.

Figure 4 - Log level net output shock- Financial Regulation

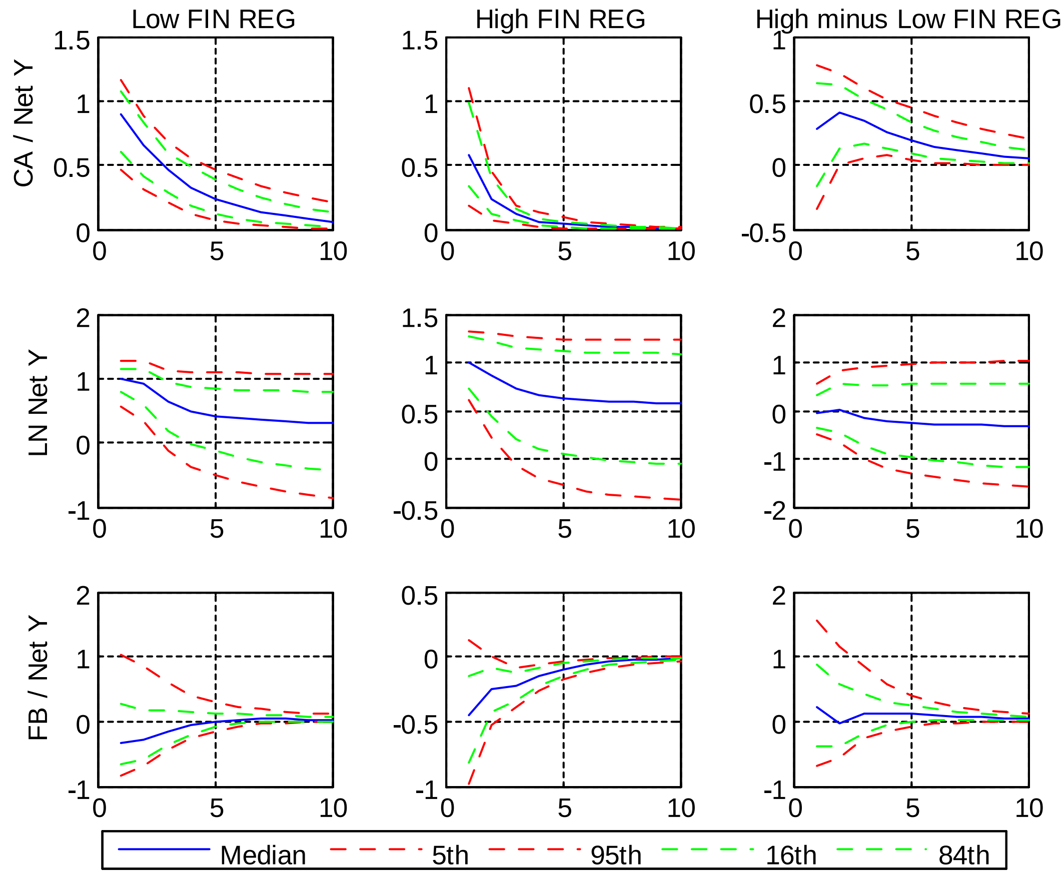


Figure 5 - Log difference net output shock-Financial Regulation

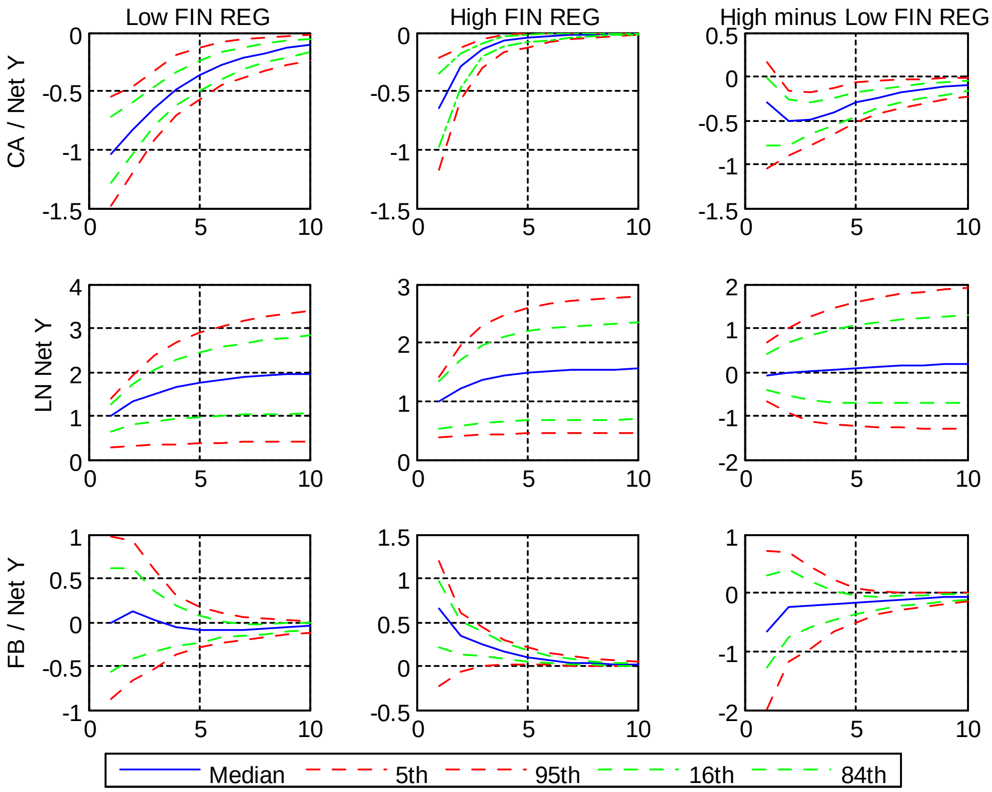


Figure 6 - Log level net output shock - Capital Account Openess

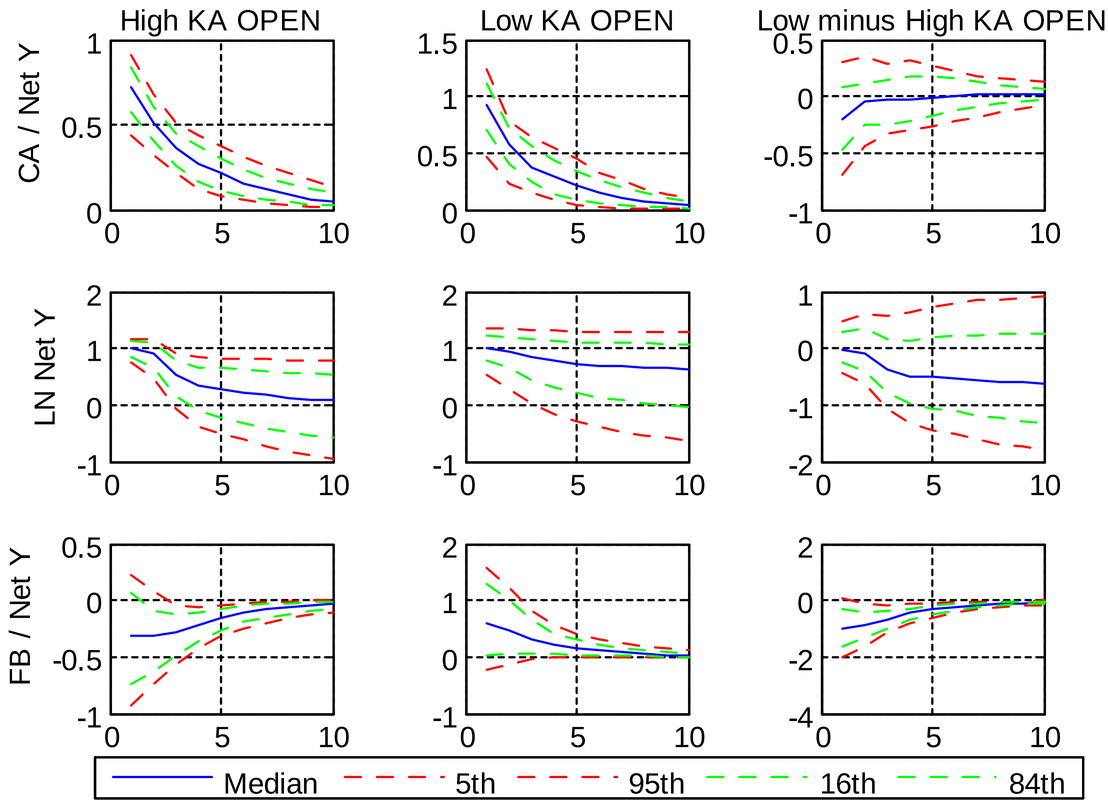


Figure 7 - Log difference net output shock - Capital Account Openess

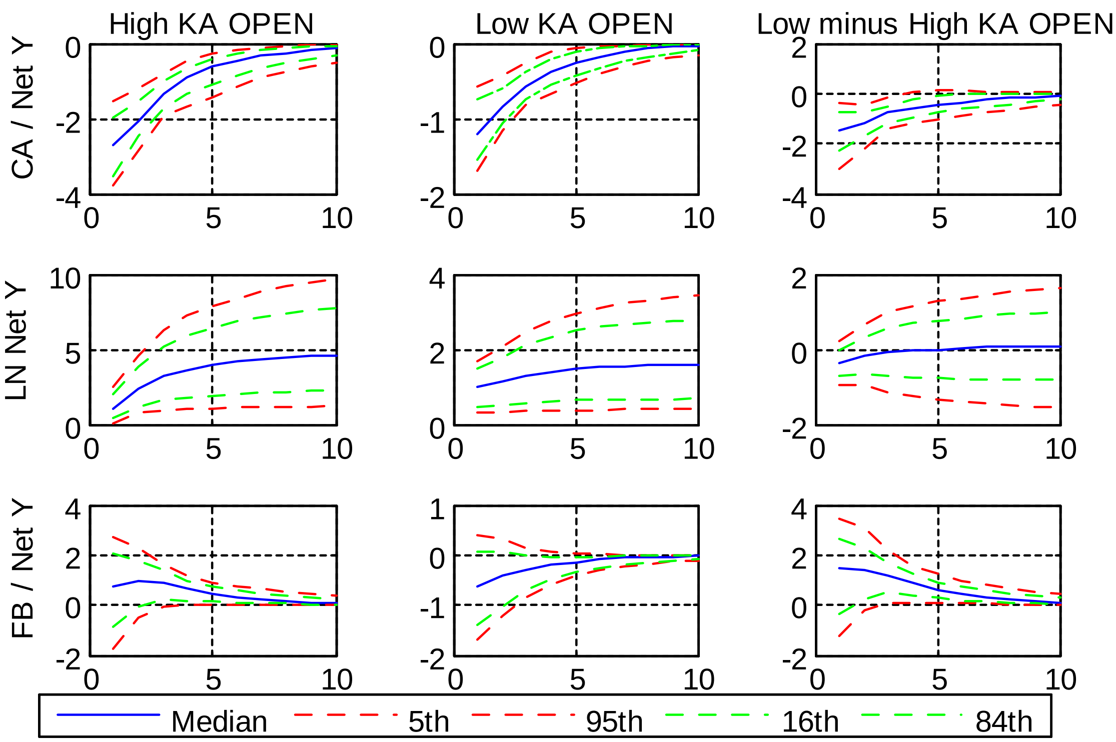


Figure 8 - Log level net output shock -Financial Regulation - Controlling for FX Regime

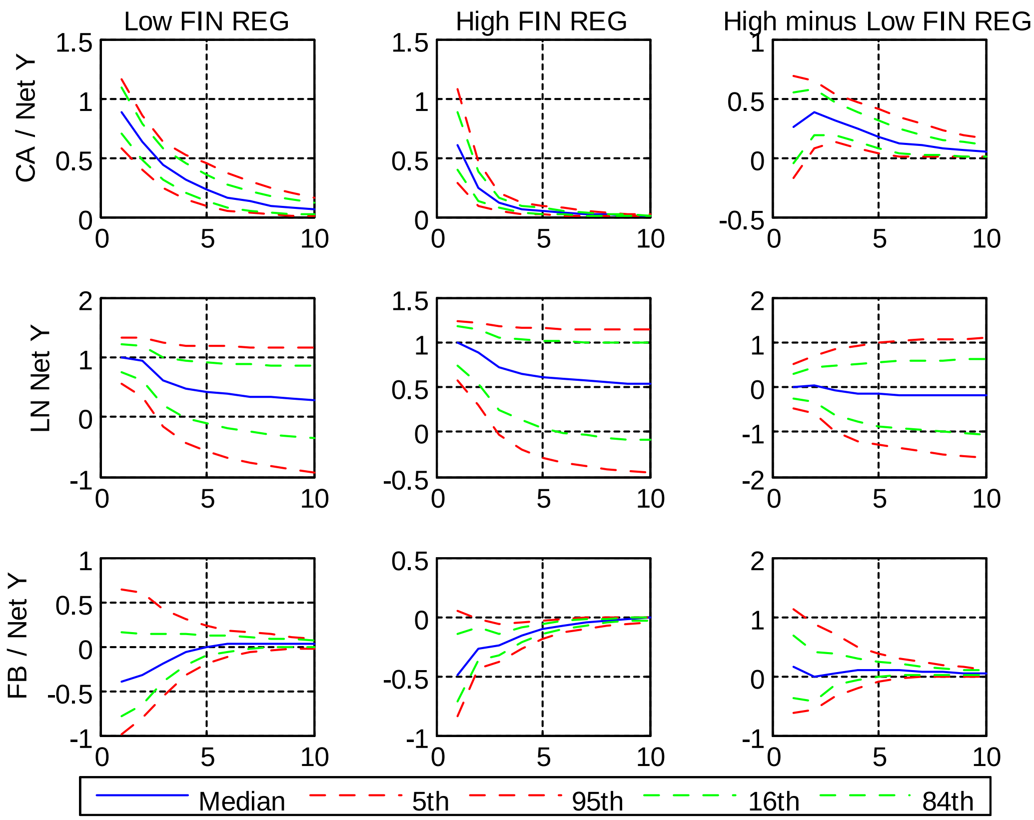


Figure 9 - Log difference net output shock - Financial Regulation - Controlling for FX Regime

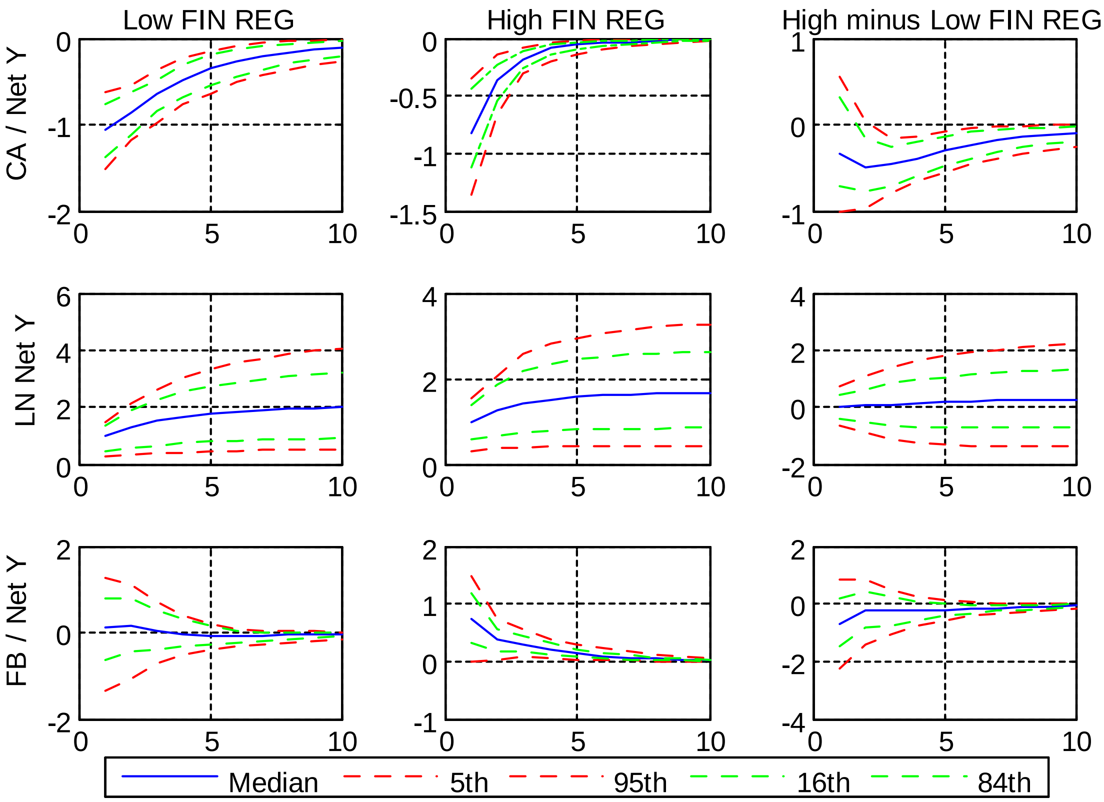


Figure 10 - Log level net output shock - Capital Account Openess - Controlling for FX Regime

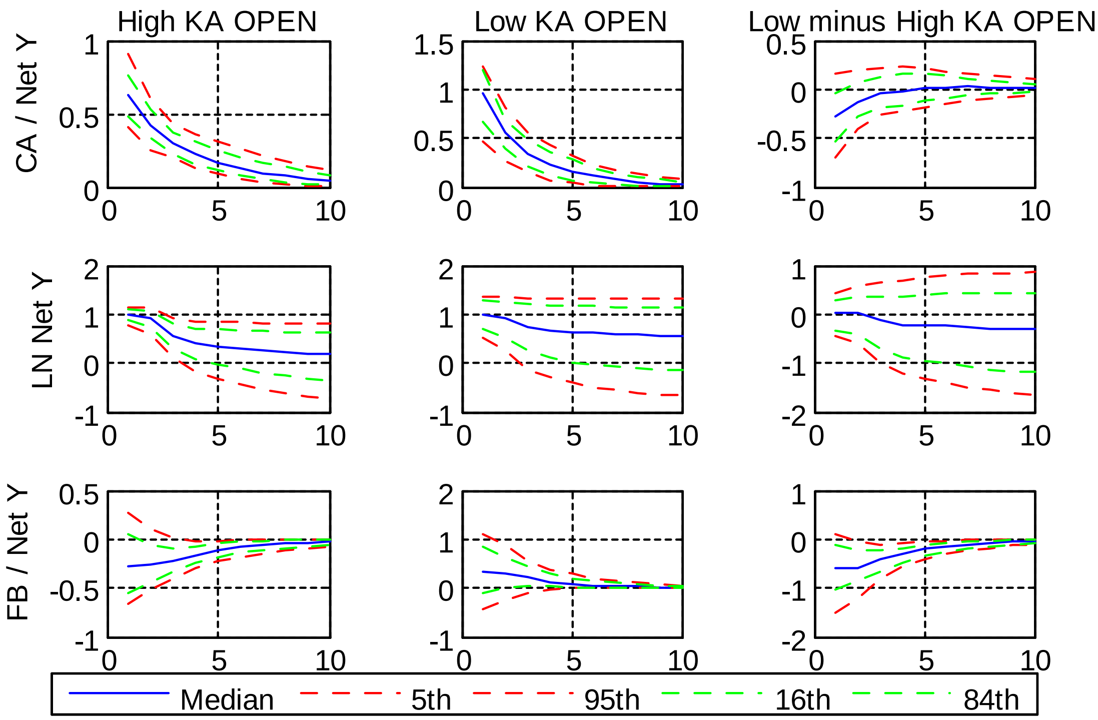


Figure 11 - Log difference net output shock - Capital Account Openess - Controlling for FX Regime

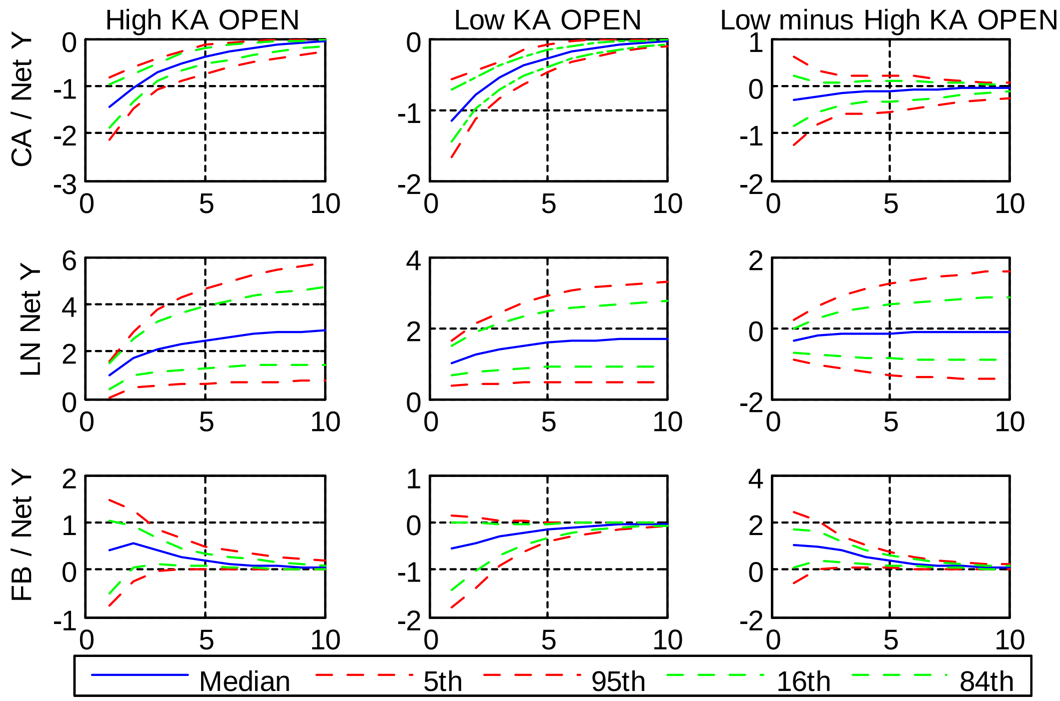


Figure 12 - Histogram of square root $\sqrt[2]{\lambda}$

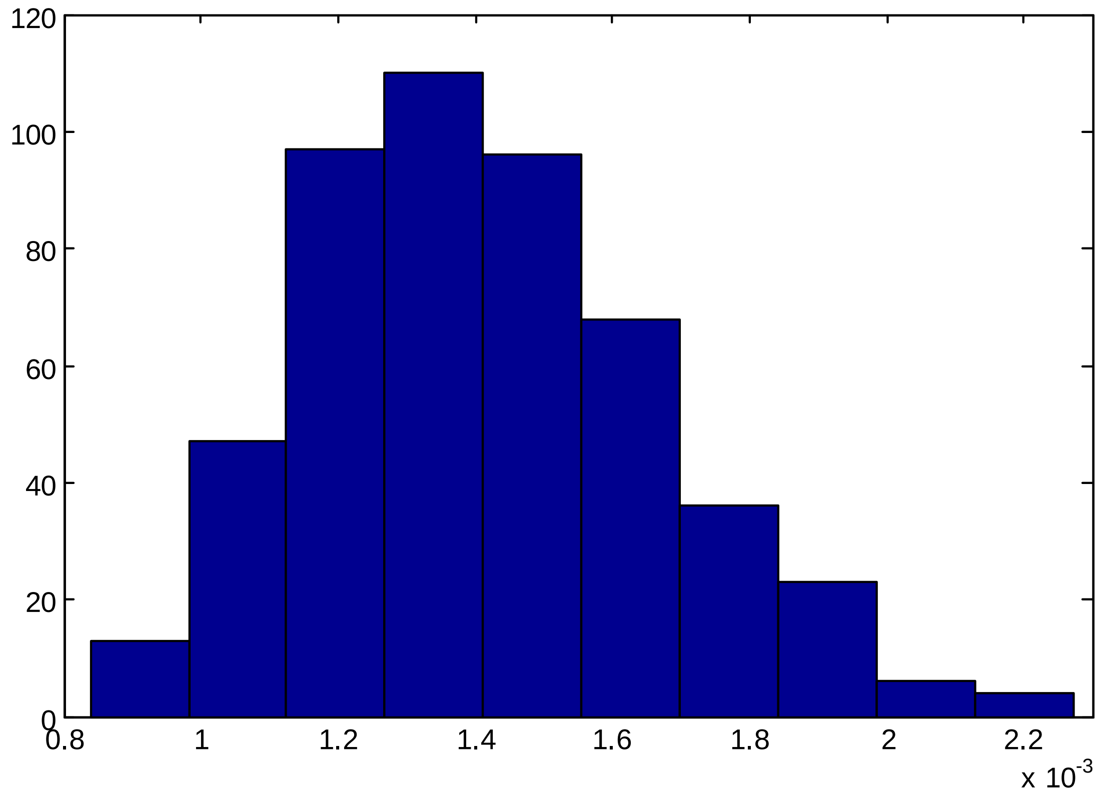


Table 2 - Countries included in the empirical analysis

Country sample		
Albania	Finland	Norway
Argentina	France	Nepal
Australia	United Kingdom	New Zealand
Austria	Georgia	Pakistan
Belgium	Ghana	Peru
Burkina Faso	Greece	Philippines
Bangladesh	Guatemala	Poland
Bulgaria	Hungary	Portugal
Belarus	Indonesia	Paraguay
Bolivia	India	Romania
Brazil	Ireland	Russian Federation
Canada	Israel	Senegal
Switzerland	Italy	Singapore
Chile	Jamaica	El Salvador
China	Jordan	Sweden
Cote d'Ivoire	Japan	Thailand
Cameroon	Kazakhstan	Tunisia
Colombia	Kenya	Turkey
Costa Rica	Kyrgyz Republic	Tanzania
Czech Republic	Korea, Rep.	Uganda
Germany	Sri Lanka	Ukraine
Denmark	Lithuania	Uruguay
Dominican Republic	Latvia	United States
Algeria	Morocco	Venezuela, RB
Ecuador	Madagascar	Vietnam
Egypt, Arab Rep.	Mexico	South Africa
Spain	Mozambique	
Estonia	Malaysia	
Ethiopia	Netherlands	