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ARTICLE

A time-varying estimation of an external reaction function for European Monetary Union countries: the role of risk-aversion and financial openness

Mariam Camarero[1], Juan Sapena[2], and Cecilio Tamarit[3]

[1]Universitat Jaume I and INTECO, Department of Economics, Castelløn de la Plana, Spain
[2]Catholic University of Valencia and INTECO, Economics Department, Valencia, Spain
[3]University of Valencia and INTECO, Department of Applied Economics II, Valencia, Spain
Corresponding author: Cecilio Tamarit; Email: cecilio.tamarit@uv.es

Abstract

This paper aims at reexamining external sustainability in a dynamic framework for nine European Monetary Union (EMU) countries during the period 1970–2021. We extend the approach of Bohn (1998) to a time-varying external reaction function. The main advantage of our empirical strategy is that it captures the dynamics of the external reaction function, by accounting for the main sources of heterogeneity among EMU countries and by including common factors like financial globalization and global risk aversion. To estimate the model, we employ a fully fledged state-space framework, which extends the simple model generally used in this literature to a panel-data time-varying parameter framework, combining fixed (common and country-specific) and varying components. Our results show an evident interplay between real and financial variables, the latter progressively increasing their importance. Although heterogeneous, the adjustment to external imbalances in most EU countries is jointly driven by the level reached in the stock of net foreign assets together with the degree of risk aversion and financial openness.

Keywords: Current account imbalances; panel unit root tests; multiple structural breaks; Kalman Filter; time-varying parameters

JEL classifications: C23; F32; F36

1. Introduction

This paper aims at reexamining external sustainability in a dynamic framework, using the database of Lane and Milesi-Ferretti (2007, 2018). We formulate empirical tests of intertemporal sustainability conditions following the approach first proposed in the seminal article of Bohn (1995) and related literature. The proposal consists of estimating a linear reaction function for the trade balance from a measure of external debt position, such as NFAs. The results obtained in previous literature on the external reaction function—see, for instance, Engel and Rogers (2006), Durdu et al. (2013), Bajo-Rubio et al. (2014)—are inconclusive concerning its existence. For this reason, we propose a more flexible estimation framework in the form of a time-varying reaction function affected by idiosyncratic (i.e., country-specific) and common factors.

In the present paper, we extend Bohn's approach in a time-varying external reaction function estimated for a panel of nine European Monetary Union (EMU) countries between 1970 and 2021, constituting a natural experiment of external adjustment. The issue of persistent external disequilibria is, from the 1990s, of relevance for many developed and developing countries but has gained momentum after the creation of the EMU. EMU membership meant eliminating the exchange rate risk for the incumbent EU countries. This brought easier conditions for borrowing in international markets due to both an increase in the supply of available funds and a decrease in cost. In other words, allowable external deficits would be higher in a monetary union as it relaxes the external intertemporal constraint. This was particularly true in the case of the peripheral countries of the EMU. However, the Great Recession (GR) was an asymmetric shock that questioned external sustainability and adjustment capacity in a monetary union. The GR meant a dramatic reassessment of external credit risk, tightening the financing conditions, and a sudden stop in capital flows, leading to an important correction in flow imbalances.[1]

We assume a lack of information in the financial markets to describe the herding behavior of lenders. Lenders mimic other investors and do not search for information by themselves. Our theoretical approach is based on Masson (1999), where the fundamentals are determined using a simple balance of payments approach. The reason for the possibility of multiple equilibria in Masson's model is higher debt service costs when devaluation (or the euro area break up in our case) is expected. This model is a benchmark for our results when capital flows and herding behavior are included. The fragility of the credit market equilibrium renders the country vulnerable to sudden changes in capital flows. Therefore, financial openness differs depending on the public's expectations. If the expected probability of bankruptcy is low, financial openness relaxes the external constraint, and, under insufficient information and moral hazard problems, it leads to excessive foreign borrowing and indebtedness. On the contrary, if the probability of bankruptcy is high, a high degree of financial openness facilitates any small adverse event to trigger a total cease of foreign financing and a balance-of-payments crisis.

The main advantage of our empirical approach is that it captures the dynamics of the external reaction function, accounting for the primary sources of heterogeneity among the EMU countries. In particular, we explore the incidence of financial globalization and global risk aversion on the model dynamics for the selected countries during the period. To estimate the model, we employ a fully fledged state-space framework that extends the simple scheme usually used in the empirical literature into a panel-data time-varying parameter setting. We estimate the state-space model through the Kalman filter (KF) and contribute to previous work in four directions. First, we modify the seminal Hamilton's GAUSS code[2] to fit a time-varying multi-parameter model; second, we extend the KF estimation from a single-country model to panel data; third, we adapt the transition equation to include control inputs according to our testing hypotheses; finally, both fixed parameters in the measurement equation and the control variables are allowed to be either common or idiosyncratic. To this end, we have developed a complete GAUSS toolbox called *sspaneltvp*.[3]

The remainder of the paper is organized as follows: Section 2 describes the theoretical model and empirical specification; Section 3 briefly reviews the previous literature; Section 4 presents the empirical methodology; and Section 5 discusses the results. Finally, Section 6 concludes.

2. The intertemporal approach to the current account and external debt sustainability

Since the seminal contributions of Obstfeld and Rogoff (1995, 1996), the intertemporal approach to the current account has been the workhorse framework in the international finance literature. This approach extends the permanent income hypothesis under rational expectations to an open economy. The current account balance is given by the sum of the following components: net exports (NX_t), net foreign income (NFI_t) due to net factor payments from abroad, and net unilateral transfers (meaning public and private transfers received from abroad minus transfers sent abroad) (UT_t).

$$CA_t = NX_t + NFI_t + UT_t \tag{1}$$

Similarly, the current account also represents the change in the value of net claims on the rest of the world and, thus, the net increase in foreign asset holdings, *NFA*. The Net Foreign Assets position, *NFA*, also named Net External Position or International Investment Position (*IIP*),[4] is the difference between the value of foreign assets owned by domestic residents, *FA*, (also denominated External Assets, *EA*), minus the value of domestic liabilities to the rest of the world, *FL*, (or also External Liabilities, *EL*). If its net foreign asset position is positive ($NFA > 0$), the country is a net creditor to the rest of the world. Conversely, if *NFA* is negative ($NFA < 0$), the country is a net debtor. In the most basic form, where the only asset is a risk-free bond earning a constant real interest rate, the current account can be written as the change in a country's net foreign asset:

$$CA_t = \Delta NFA_t = NFA_t - NFA_{t-1} \tag{2}$$

Therefore, future current account and net foreign asset positions are related to the current account and net foreign asset positions through future net foreign income flows[5]

Following Gourinchas and Rey (2007), we start from a country's intertemporal budget constraint and derive two implications. The external constraint implies that today's imbalances must predict either future changes in the trade balance (flow adjustment), future movements in the returns of the NFA portfolio (changes in the stock of foreign assets), or both. Due to the heterogeneity of frictional information between financial and real goods markets, most of the adjustment goes through asset returns in the short and medium term. In contrast, at longer horizons, it occurs via the trade balance. Therefore, if the current account is in deficit ($CA < 0$), the change in the net foreign asset position is negative, indicating that the increase in foreign debt was more significant than the increase in foreign assets over the year. A negative change in the net foreign asset position is referred to as a net capital inflow since more capital flowed into the country through additions to the level of foreign debt than flowed out through purchases of foreign assets.

An economy is deemed externally solvent[6] if the intertemporal external budget constraint (IEBC) is satisfied. The IEBC stipulates that the present value of future trade surpluses should be equal to current net external liabilities. Moreover, according to the definition of Milesi-Ferretti and Razin (1996), external positions are deemed unsustainable if the fulfillment of the IEBC is likely to require a drastic adjustment in macroeconomic fundamentals (e.g., severe demand compression or currency devaluation) or an outright external crisis.[7]

To derive the different testing hypotheses, consider the accumulation identity for net foreign assets between t and $t-1$:

$$NFA_t = A_{t-1} \cdot (1+i_A) \cdot \frac{E_{t-1}}{E_t} - L_{t-1} \cdot (1+i_L) \cdot \frac{E_{t-1}}{E_t} + CA_t \tag{3}$$

In (3), a nation's international investment position therefore depends on three elements: the interest rates on international assets A and liabilities L, namely, i_A and i_L, the rate of change in the exchange rate $\frac{E_{t-1}}{E_t}$, and the current account balance CA_t.

Several simplifying assumptions about these elements are needed to keep the model basic. First, we assume a world without capital market frictions, so that $i_A = i_L$;[8] second, we assume a world of stable exchange rates $\frac{E_{t-1}}{E_t}$;[9] third, there are no unilateral transfers. The rates of return on foreign assets and liabilities influence the extent of these flows. Net foreign income is essentially the difference between interest earned on foreign assets and interest paid on foreign liabilities. Under this assumption, provided our first assumption about frictionless capital markets, we know that apart from the trade balance, the only other non-zero item in the CA is the net foreign income and assume that the primary bulk of the current account comes from the trade balance $CA_t = NX_t$, then we get the expression:

$$NFA_t = (1+i_t) \cdot NFA_{t-1} + NX_t. \tag{4}$$

These variables can be expressed in nominal terms, in real terms, or, as in Chortareas et al. (2004), as a ratio to GDP, as long as i is adjusted accordingly. If the variables are in nominal terms, r represents the nominal interest rate; when the variables are in real terms, r is the real interest rate; and finally, if the variables are defined as ratios to GDP, r_t is the growth-adjusted real interest rate that follows from dividing the gross real interest rate by the gross rate of output growth, $r_t = \frac{(1+i_t)}{(1+g_t)} - 1$. We will refer to the growth-adjusted real interest rate as r_t. If we assume the GDP growth to be constant, g, then $r = \frac{(i-g)}{(1+g)}$. In this last case, both CA and NFA are defined as ratios to GDP, and the notation changes for:

$$nfa_t = (1 + r) \cdot nfa_{t-1} + nx_t. \tag{5}$$

The equations displayed so far characterize the period-by-period balance of payments. They equate the current account deficit (surplus) to capital inflow (outflow) or net borrowing (lending) from (to) abroad. Iterating (5) forward and assuming that the expected value $E(r_t|\varphi_{t-1}) = r$, with φ_{t-1} being the information set available in $t - 1$, we get:

$$nfa_t = \sum_{j=0}^{\infty} \left(\frac{1}{1+r} \right)^j E(nx_{t+j} \mid \varphi_{t-1}) + \lim_{T \to \infty} \left(\frac{1}{1+r} \right)^T E(nfa_{t+T} \mid \varphi_{t-1}). \tag{6}$$

Equation (6) states that international agents can lend to an economy if they expect that the present value of the future stream of net export surpluses equals the current stock of net foreign assets. Hence, the sustainability hypothesis, or long-run external budget constraint, implies that:

$$\lim_{t \to \infty} \left(\frac{1}{1+r} \right)^T E(nfa_{t+T} \mid \varphi_{t-1}) = 0. \tag{7}$$

This transversality condition means that the present value of the expected nfa stock when t tends to infinity must equal zero, that is, a no-Ponzi game condition. If equation (7) holds, (6) means that sustainability is possible with current account deficits as long as they do not grow faster than the output of the expected value ($E(g_t|\varphi_{t-1}) = g$). In practical terms, solvency arithmetic examines whether the net external position/GDP ratio (nfa) grows more or less rapidly than the difference between the real interest rate (r) and the growth rate of the economy (g).

In other words, the external net asset position is sustainable if the stock today is less than or equal to the present value of net exports. In a deterministic setting with non-stationary data, these conditions have been empirically interpreted as requiring cointegration between imports and exports with a cointegrating vector $\left(1, -\beta_{1,0}\right)$, where $\beta_{1,0} \leq 1$.[10]

In a stochastic environment, GDP is affected by multiple shocks of different nature. The sustainable intertemporal external constraint will hold if a long-run positive (negative) relationship exists between exports (imports) and the level of the net external asset position. This implies an endogenous policy response of the balance of payments to maintain the net external position at some constant optimal ratio to income (nfa*).

For this equilibrium, exports (imports) must respond positively (negatively) to deviations from the net external position relative to the optimal ratio. As Bartolini and Cottarelli (1994) demonstrate, external solvency requires that the asymptotic growth rate of the economy exceeds the asymptotic interest rate. However, since the long run takes considerable time to materialize, policymakers may introduce a reaction function as an endogenous policy response. This helps governments avoid episodes where negative nfa exceeds certain threshold values by stabilizing it around a target level (nfa*).

Importantly, this ratio nfa* need not be strictly "optimal." Rather, exports (imports) can respond positively (negatively) to nfa deviations beyond the optimal ratio. Consequently:

- Countries with steady-state creditor positions (NFA $>$ 0) should maintain trade deficits (NX $<$ 0)
- Countries with steady-state debtor positions (NFA $<$ 0) should maintain trade surpluses (NX $>$ 0)

When $g > r$, the economy can sustain persistent trade deficits (importing more than exporting) without the net external position exceeding the desired level (nfa*). However, if the growth rate falls below the real interest rate ($g < r$), the discount rate decreases, necessitating corrective measures. These measures promote current account surpluses to reduce future debt burdens and stabilize nfa* over time.

This means a long-run stock-flow relationship between the CA flow variables and the stock of the NFA will be present. However, this basic definition suffers from several shortcomings as highlighted by Bohn (2007), that questioned the concept of sustainability and its application in the context of unit roots and cointegration.[11] More specifically, he shows that any finite order of integration of these series separately also leads to the fulfillment of the IEBC. Moreover, the cointegration relationship between exports and imports is unnecessary for the no-Ponzi game condition to hold. However, a long-run relationship following an error-correction specification between nfa_t and nx_t should exist to avoid any explosive outcome from the variables that determine the external equilibrium in the long run. We propose an alternative approach to test for external sustainability based on an external reaction function.

3. An external sector reaction function: The Bohn's approach

External sustainability analysis has evolved along several methodological paths. Early research concentrated on two main approaches: current account stationarity[12] and external debt stock analysis.[13] A third approach emerged through cointegration testing of long-run relationships between exports and imports.[14]

The relationship between current account (CA) and net foreign asset position (NFA) has received less attention. Wickens and Uctum (1993) developed a two-equation VAR with endogenous primary deficit, arguing that negative feedback from net national indebtedness to trade deficit suffices for IEBC satisfaction. Ahmed and Rogers (1995) studied cointegrating vector stability during unusual events like wars. Engel and Rogers (2006) applied Bohn's approach to US external balance sustainability, while Durdu et al. (2013) adapted Bohn (2007)'s error-correction reaction function to panel framework, finding significant negative NX-to-NFA response. Camarero et al. (2013) employed multicointegration for IEBC testing.

Recent studies have incorporated NFA dynamics using nonlinear error correction models. This empirical work extends IEBC analysis to stochastic settings (Bohn, 1998). Bohn (1995, 1988) demonstrated external sustainability's dependence on external debt discount rate changes ("valuation effects"), suggesting traditional cointegration tests' oversimplification.

Bohn (1998, 2007) notes IEBC fulfillment imposes weak econometric restrictions. Extending his fiscal framework to external constraints shows IEBC satisfaction requires either *nfa* or *nx* series stationarity after finite differencing. Bohn (2007) proposes a policy reaction function approach[15] as cointegrating conditions are merely "sufficient" for transversality. This reaction function, initially conceived for public finances imbalances, assumed a time-invariant relationship between the primary fiscal deficit and the debt ratio. By analogy, it would also posit a time-invariant relation between what we can call the primary external balance[16] and the external financial stock ratio, which seems at odds with a simple visual data inspection. This can be observed in Figure 1, which shows a relationship that appears to vary over time. Thus, in this paper, we modify the original parametric specification proposed by Bohn (1998) to allow for a time-varying relationship estimated through the Kalman filter.

Bohn (2007) proposes an alternative approach for sustainability testing based on an arbitrarily high order of integration of the variables involved and error-correction model-type policy reaction functions. Bohn's ECM approach decouples the concepts of solvency and the external intertemporal budget constraint. According to Bohn, it is sufficient for the transversality condition (7) to hold for a country that its NX and NFA over GDP ratios (*nx* and *nfa*) follow an error-correction specification of the form:

$$nx_t - \delta nfa_{t-1} = z_t \sim I(m) \tag{8}$$

where the parameter $\delta < 0$ is such that $|\delta| \in (0, 1 + r)$.

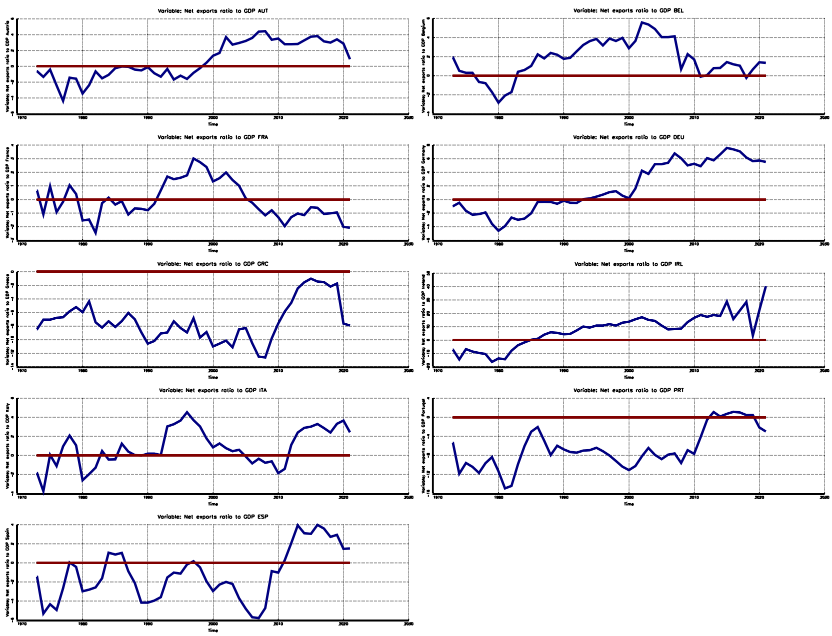

Figure 1. Net exports ratio to GDP. Selected EMU countries. 1971–2015.

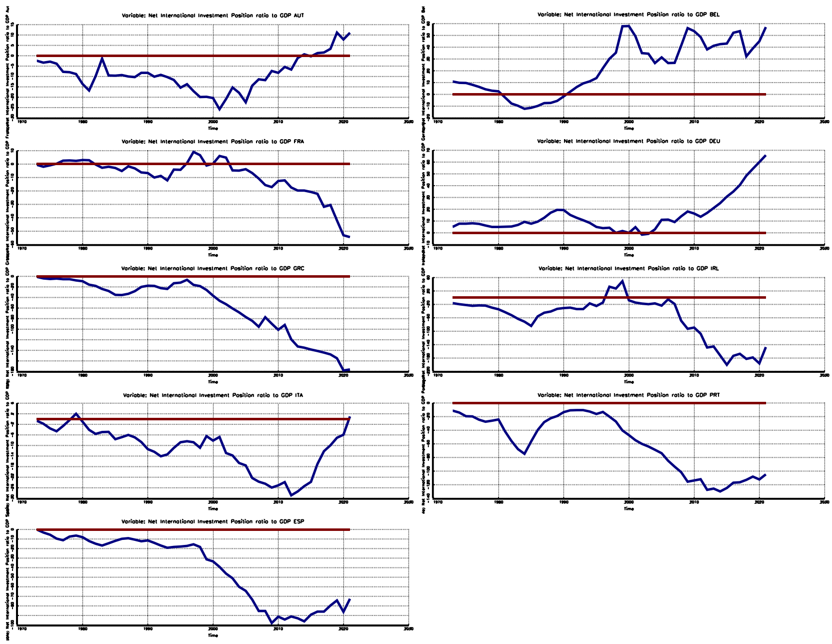

Figure 2. International investment position ratio to GDP. Selected EMU countries. 1973–2021.

Reaction functions with $|\delta| > r$ imply stationary *nfa* stocks and *nx* imbalances. $0 < |\delta| < r$ imply mildly explosive paths for external debt and *nx* deficits, but growing slowly enough to be consistent with the IEBC. Reaction functions with a $|\delta| = r$ constitute a particular case that implies a difference-stationary *nfa* and stationary *nx*; hence it satisfies the unit root and cointegration conditions.

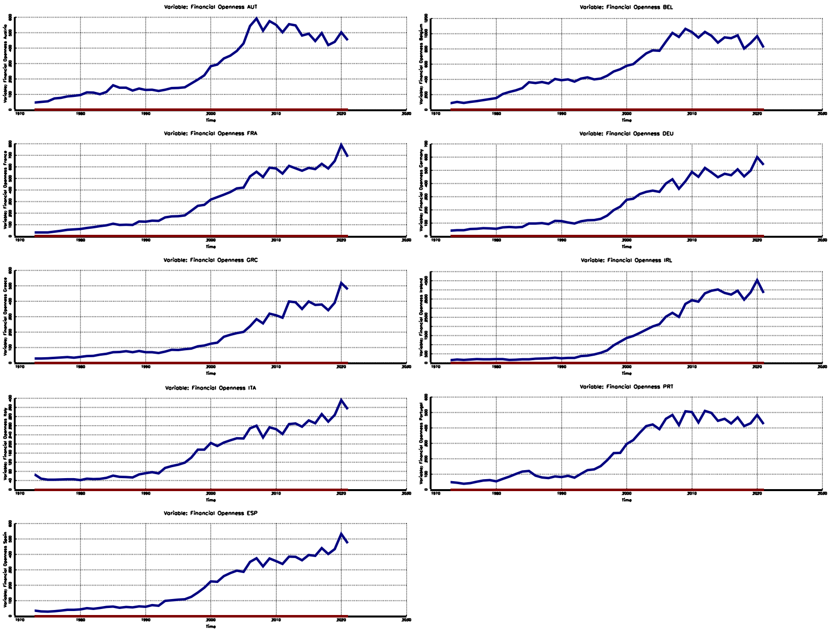

Figure 3. Financial openness. Selected EMU countries. 1971–2015.

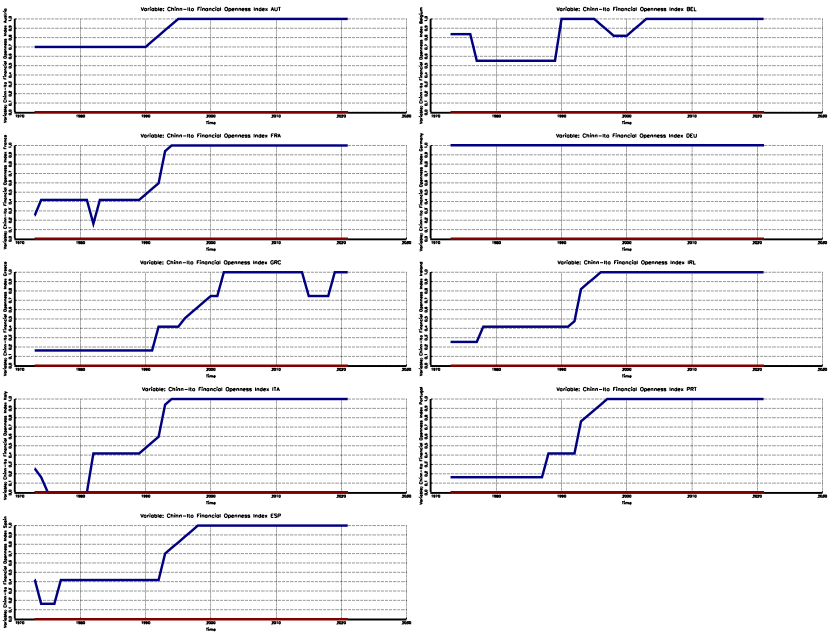

Figure 4. Chinn-Ito financial openness index. Selected EMU countries. 1971–2015.

The economic intuition is that the transversality condition holds if a debtor responds positively to debt accumulation through $\delta < 0$. Such a reaction function implies that economic agents (both the private sector and the government) adjust their savings and investment plans over time in line with the financing requirements implied by changes in the economy's *nfa* position. The above proposition constitutes only a sufficiency condition for the IEBC to hold. Thus, it is possible that

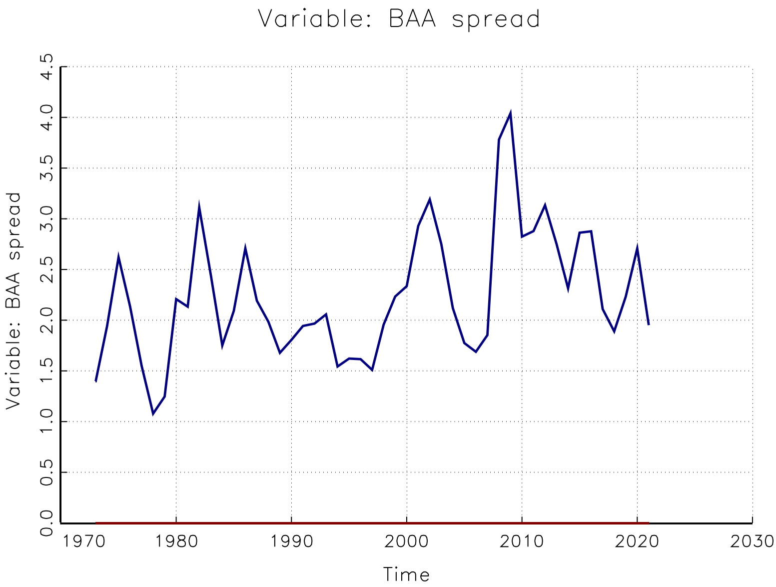

Figure 5. BAA spread. 1971–2015.

countries might not display a negative response coefficient, with *nfa* stationary of finite order, so that IEBC holds.

As in Trehan and Walsh (1991), the critical technical condition is that all roots are strictly less than $1 + r$, and this condition is not affected by allowing for unit roots. Practical challenges should be acknowledged at this point. Because the appropriate discount rate is typically a small number, it may be difficult to distinguish empirically between a unit root and a $(1 + r)$-root in the debt process and between values for δ that are slightly greater or slightly less than zero.

In the present paper, we develop an error correction model approach to test for the external sustainability of nine euro area countries. We use annual data from 1970 to 2021 on trade balance (NX), gross national product, and the stock of Net Foreign Assets (NFA). In addition, the model includes some control variables. As Bohn (2007) highlights, assessing sustainability is not a mechanical exercise, as the researcher must face several specification issues, such as defining the appropriate control variables and discount rates (notably if debt management is endogenized). To identify the marginal effect of *nfa* on *nx* in the external reaction function, theories of optimal consumption smoothing suggest that one must control for temporary shocks to government outlays and income, as these trigger higher than-normal imbalances via the absorption approach. Therefore, both variables should enter with a negative sign.

We claim external sector reaction functions can provide a more robust framework for assessing external sustainability. The analysis should examine how agents (both public and private) adjust *nx* in response to changes in *nfa*. Following Bohn's approach, we posit that external position sustainability can be maintained even in an uncertain environment if *nx* responds positively to increases in *nfa* stock. This can be tested by examining whether $\beta < 0$ in:

$$nx_t = \beta \cdot nfa_{t-1} + \delta \cdot Z_t + \varepsilon_t = \beta \cdot nfa_{t-1} + \mu_t \tag{9}$$

where Z_t represents a vector of primary external surplus determinants operating through parameter δ, and ε_t denotes the error term.

The intuition behind an external reaction function, specifically, how net exports react to lagged net foreign assets accumulation, warrants clarification. Unlike fiscal reaction functions, where governments adjust the primary surplus to service debt, external reaction functions involve multiple institutional sectors (public and private, financial and non-financial). In a stochastic environment, domestic debtors must signal their commitment to external debt repayment by improving their external primary surplus, approximated by the net exports ratio. This adjustment manifests differently across countries. The optimal path occurs through returns on productivity gains from capital accumulation, though some countries may require nominal depreciation to enhance competitiveness. In monetary unions, this improvement often necessitates a challenging process of real depreciation. The external constraint remains binding in all cases, even when default becomes necessary.

Additionally, Bohn (2007) suggests the possibility of a time-varying setup. Within this framework, if both the debt stock and the primary surplus are nonstationary, while μ_t is stationary, equation (9) is equivalent to the cointegration test suggested by Trehan and Walsh (1991). However, if they are stationary, we need to consider potential determinants of the primary surplus to avoid biased coefficient estimates. To circumvent the problem of dependence on these two variables for the fiscal reaction function, Bohn includes a business cycle variable, *YVAR*, and also a measure of the deviation of actual public expenditure from its long-run trend, *GVAR*, based on the tax-smoothing theory of Barro (1979, 1986) and other related literature.

In the present paper, we follow the suggestion made in Bohn (2007)[17] to transpose the fiscal reaction framework to a external reaction function, where B (the agent's net debt) represents now a country's net external liabilities, T (the agent's income) corresponds to exports, G (agents' outlays) is now redefined as imports and r stands for the interest rate on external liabilities. This new framework can be an interesting contribution since if a country's external sector reacts to an increase in its net external position by adjusting its net export surplus, it signals to the markets its ability (and willingness) to restore the external stance to a sustainable path. According to Bohn (1995, 2007) this dynamic equilibrium represents an error correction mechanism: if *nfa* increases, agents (public and private) should respond by improving the primary balance[18] (*nx*), to offset and even reverse the rise in the *nfa* ratio.

A salient feature of the present paper is that as in Marinkov and Burger (2012) and related documents, we test for a time-varying external reaction function, represented in state space form and estimated using the Kalman filter algorithm, where the varying parameter vector of net exports to net foreign assets ratio, follows a partial adjustment mechanism mean reverting (MRV) model as follows:

$$\beta_{i,t} = \Phi\beta_{i,t-1} + (1-\Phi)\,\bar{\beta}_i + \nu_{i,t} \tag{10}$$

where ν_{it} is a Gaussian error with a zero mean and a fixed variance, gradually making the parameter return to its mean.

The mean-reversion model represents a general modelization of the parameters: the OLS model is obtained when $var(\nu_{i,t}) = 0$; when $\Phi = 1$ we obtain a random walk model for the varying parameters; and when $\Phi = 0$ we have a random coefficient model where the coefficient fluctuates randomly around a mean value.

The MRV model can be rewritten as:

$$\left(\beta_{i,t} - \bar{\beta}_i\right) = \Phi\left(\beta_{i,t-1} - \bar{\beta}_i\right) + \nu_{i,t} \tag{11}$$

Time-varying parameter (TVP) regression models are an interesting application of state-space models, where the unobserved component, $\xi_{i,t} = \left(\beta_{i,t} - \bar{\beta}_i\right)$, evolves along time according to the expression:

$$\left(\beta_{i,t+1} - \bar{\beta}_i\right) = F\left(\beta_{i,t} - \bar{\beta}_i\right) + \upsilon_{i,t+1} \tag{12}$$

As stated in Hamilton (1994b), assuming that the eigenvalues of F are all inside the unit circle, the fixed coefficient, $\bar{\beta}_i$, can be interpreted as the average or steady-state coefficient vector, while $\xi_{i,t} = \left(\beta_{i,t} - \bar{\beta}_i\right)$ is the varying deviation from its mean parameter, represented here by the TVP unobserved parameter $\xi_{i,t}$.

One way to motivate such an adjustment mechanism could be as follows. First, let us assume that a country may have a desired level of adjustment in net exports (*nx*) in reaction to changes in *nfa*. Define this level as $\widetilde{\beta}_{i,t}$. Here, we can suppose that such a reaction is conditioned by several factors: degree of financial openness, global risk aversion, economic juncture (whether GDP growth is a boom or recession regime), etc. All these variables are gathered in a vector $x_{i,t}$. Also, some unobserved variables influence the countries' reaction. Assume that they are captured by a random variable $\nu_{i,t}$. We can therefore write:

$$\widetilde{\beta}_{i,t} = \alpha_0 + \alpha_1 \cdot x_{i,t} + u_{i,t} \tag{13}$$

Then, we can assume that it takes time for the government to reach the desired level of adjustment and suppose, for instance, that the adjustment is partial and revised regularly proportionally to the discrepancy from the target. Let us consider, for instance:

$$\beta_{i,t} - \beta_{i,t-1} = \lambda \cdot (\widetilde{\beta}_{i,t} - \beta_{i,t-1}) \tag{14}$$

This equation takes the form of a Mean-Reverting (MRV) model

Combining both equations yields:

$$\beta_{i,t} = \alpha_0 \cdot \lambda + (1 - \lambda) \cdot \beta_{i,t-1} + \lambda \cdot \alpha_1 \cdot x_{i,t} + \lambda \cdot u_{i,t} \tag{15}$$

or

$$\beta_{i,t} = \mu_0 + \Phi \cdot \beta_{i,t-1} + 1 - \Phi \cdot \alpha_1 \cdot x_{i,t} + \nu_{i,t} \tag{16}$$

where $\mu_0 = \alpha_0 \cdot \lambda$, $\Phi = 1 - \lambda$, and $\nu_{i,t} = \lambda \cdot u_{i,t}$. If we consider $\nu_{i,t} = 0$

The fact that a well-known partial adjustment model leads to an MRV adjustment mechanism has several advantages. First, instead of considering the variables in a way that may seem ad hoc, we can justify our formal approach using a standard adjustment mechanism commonly used in economic modelization. Second, this representation also has empirical advantages.

The above choice made on specification permits the joint inclusion of both fixed ($\bar{\beta}$) and varying (ξ_t) components for the parameters in the measurement equation[19] as follows:

$$y_t = x_t^{'}\bar{\beta} + x_t^{'}\xi_t + \omega_t \tag{17}$$

Using the above framework, we estimate a time-varying external sector reaction function for the euro area countries in a modelization where, inspired by Barro's tax-smoothing model (Barro, 1979), we include other non-debt determinants of the primary surplus as control variables. The choice of this variable is not trivial in terms of model specification. In this paper, as in Mendoza and Ostry (2008), we include a business cycle variable (YVAR) as well as the estimated level of temporary imports over the cycle (MVAR). YVAR and MVAR are detrended GDP and imports, respectively. These variables are frequently obtained using the Hodrick-Prescott filter.[20] However, in this paper, we use the filter recommended by Hamilton (2018), who proposes an alternative concept of the cyclical component of a possible nonstationary series with better properties.

Following Barro (1986), YVAR and MVAR are calculated as:

$$YVAR_t = \frac{(y_t^* - y_t)}{y_t^*} \cdot \frac{m_t^*}{y_t} \tag{18}$$

$$MVAR_t = \frac{(m_t - m_t^*)}{m_t^*} \cdot \frac{m_t^*}{y_t} \tag{19}$$

where m_t and y_t are, respectively, the imports and income levels at time t, while m_t^* and y_t^* are their trend levels.

Concerning the expected signs of the variables, *nfa* should have a negative coefficient (as it is defined as assets minus liabilities, a positive *nfa* position implies a net creditor position). As the Barro model predicts, the sign of YVAR is expected to be negative, implying a counter-cyclical policy behavior. Finally, MVAR also enters negatively. The inclusion of a measure of cyclical (temporary) imports isolates the cyclical pattern in the imports-to-income elasticity from its trend (Bussière et al. 2013), normally driven by the evolution of demand components (that typically have a high import content), such as investment.

The empirical approach minimizes the endogeneity problems likely to appear in estimating reaction functions. Indeed, the output gap may be somewhat correlated with the primary balance (*nx*) through an absorption effect. In contrast, the *nfa* could be correlated with the residuals, creating a downward bias on the estimated coefficient of *nfa*. However, this possible endogeneity bias tends to be minimized in an appropriate state-space representation of the model (Kim and Kim, 2011; Swamy et al. 2017).

Finally, we allow parameter β in our external reaction function to be time-varying. In the context of an external reaction function, such a time-varying policy rule is sustainable provided that $\beta_{i,t} = \bar{\beta}_i + \xi_{i,t}$ is always negative. However, as Greiner and Fincke (2015) remarked for the case of a fiscal reaction function, this condition can be too restrictive. They conclude that a sufficient condition for external sustainability is for the reaction coefficient to be negative on average. In our framework, we consider that the external sustainability is fulfilled if the mean of the fixed-parameter component is negative and significant.[21]

4. A tvp external reaction function specification

Our model captures the dynamic reaction of our group of countries to an increase in external liabilities by adjusting their net export balance. This will capture the market's ability (and the policymaker's willingness) to restore the external stance to a sustainable path.

If the *nfa* stock increases, the government should respond by improving the net export balance. A usual specification of an external reaction function for annual data would show the reaction of *xn* to changes in one-period lagged *nfa*, controlling for other variables.

In the present exercise, the time-varying parameters general specification of the external reaction function is represented by the equation:

$$\begin{aligned} xn_{i,t} = \beta_{0i} + \beta_{1i} xn_{i,t-1} + \beta_{2i} nfa_{i,t-1} + \beta_{3i} nfac_{i,t-1} + \gamma_{1i} yvar_{i,t-1} + \gamma_{2i} mvar_{i,t-1} \\ + \xi_{i,t} nfa_{i,t-1} + \omega_{i,t} \end{aligned} \tag{20}$$

where $nfac_{i,t-1} = nfa_{i,t-1} \cdot \ln(finopen_{i,t-1})^{-1}$.

An alternative rearrangement of equation (20) is given by:

$$\begin{aligned} xn_{i,t} = \beta_{0i} + \beta_{1i} xn_{i,t-1} + \left(\beta_{2i} + \beta_{3i} \ln(finopen_{i,t-1})^{-1} + \xi_{i,t}\right) nfa_{i,t-1} + \gamma_{1i} yvar_{i,t-1} \\ + \gamma_{2i} mvar_{i,t-1} + \omega_{i,t} \end{aligned} \tag{21}$$

Hence, in our model, the dependent variable is $xn_{i,t}$, which stands for the net exports ratio to GDP, while the explanatory (country-specific parameter) variables are the intercept, the lagged dependent, as well as $nfa_{i,t-1}$, that stands for the lagged value of the net foreign assets-to-GDP ratio. In addition, we include $nfac_{i,t-1}$ to account for the relative financial openness of the country, denoted $finopen_{i,t}$. Using a similar approach to Kraay (1998), our *de facto* measure of financial openness is obtained as the sum of external assets and liabilities as a share of GDP. As the expected relationship is non-linear, we take the natural log of the indicator; moreover, as our intuition is that more profound financial openness contributes to diluting the relative importance of financial positions, we use the inverse of our measure of financial openness, $1/\ln(finopen_{i,t})$.

The variables $yvar_{i,t-1}$ and $mvar_{i,t-1}$ are measures of cyclical variations in output and temporary imports over the cycle, calculated following Hamilton (2018) approach.[22] Finally, $\xi_{i,t}$ stands for the time-varying parameter for $nfa_{i,t-1}$.

The state-space representation of the model also includes the transition of the unobserved elements following the partial adjustment described in the previous section. Each of the elements of the state vector, $\xi_{i,t}$, follows a stochastic process such as:

$$\xi_{i,t+1} = \phi\xi_{i,t} + \mu_{1i}[baas_t * kaopen_{i,t}] + \mu_{2i}kaopen_{i,t} + \upsilon_{i,t+1} \quad (22)$$

where $\upsilon_t \sim N(0, Q)$, and the transition of the unobserved vector is driven by an autoregressive component with parameter ϕ. Additionally, two control instruments are expected to govern this transition: global risk aversion, measured by the BAA spread, $baas_t$ and the degree of financial openness, measured by the Chinn and Ito (2007) index, $kaopen_{i,t}$. These control inputs impact through their potentially country-specific parameters, μ_{1i} and μ_{2i}.

The above general state-space model can be estimated by maximum likelihood using the Kalman filter algorithm. In the next section, we will test for alternative specifications of the model embedded in this general representation related to the country-specific versus common-for-the-panel nature of estimated parameters.

Using the information contained in the IMF Annual Report on Exchange Arrangements and Exchange Restrictions, the Chinn-Ito index constitutes a popular *de jure* measure of financial openness based on the extent and intensity of capital controls. As remarked in Fratzscher and Bussiere (2004), *de jure* measures of financial openness capture different patterns compared to *de facto* ones, as a country that is open de jure may not necessarily experience such inflows.[23]

The interplay between the two variables (financial openness and risk) helps us to explain the higher debt service costs that trigger the reaction dynamics from an equilibrium of increasing external imbalances to a new equilibrium with a reduction in external imbalance. The range of country fundamentals for possible crises is more comprehensive, as capital markets are more integrated and information remains constrained.

5. Data and empirical results

5.1 The data

We analyze the evolution of the main variables involved in the reaction function of nine EMU countries (Austria, Belgium, France, Germany, Greece, Ireland, Italy, Portugal, and Spain). The data are depicted in Figures 1 to 5 and come from different sources covering the period 1970–2021.[24]

The data of exports and imports of goods and services as a percentage of GDP (with codes NE.EXP.GNFS.ZS and NE.IMP.GNFS.ZS) have been extracted from the World Bank national accounts database and the OECD National Accounts data files. GDP and its deflator have also been downloaded from the World Bank National accounts (code NY.GDP.MKTP.KD). The Net International Investment positions ratio to GDP has been obtained from the updated and extended version of the External Wealth of Nations Mark II database of Lane and Milesi-Ferretti (2007), which includes valuation effects. This is also the data source on external assets and liabilities to compute the financial openness indicator. Moreover, as a measure of the degree of capital account openness, we use the Chinn-Ito index ($kaopen_{i,t}$).[25]

Finally, as a measure of global risk aversion, we use the (natural logarithm) of the yield spread between low-grade U.S. corporate bonds (BAA) and the 10-year Treasury bonds ($baas_t$,). This is an empirical proxy for the overall investors' risk attitude. This variable has been obtained from the Federal Reserve Economic Database, provided by the Federal Reserve Bank of St. Louis.[26]

5.2 Testing for panel unit roots and structural breaks

Before estimating the TVP model, in this section, we study the order of integration of the variables, in particular, $nfa_{i,t}$ and $nx_{i,t}$, both expressed as ratios to GDP.

Table 1. Bai and Carrion-i-Silvestre (2009) structural breaks, (1973–2021)

	$xnr_{i,t}$	$nfa_{i,t}$	$nfac_{i,t}$	$finopen_{i,t}$	$kaopen_{i,t}$	n. observations
Austria						49
Belgium						49
France						49
Germany		2012	2012			49
Greece		2000	2000			49
Ireland					1981	49
					1992	
				1986		
Italy		2014	2013	1999		49
				2011		
		1985	1985			
Portugal		1992	1994			49
		2015	2013			
Spain		1999	1999			49
		2010	2010			

Note: Bai and Carrion-i-Silvestre (2009) estimations, allowing for up to 3 breaks.

The univariate properties of the data determine the choice of the most suitable specification of the model. Therefore, we also consider the potential existence of unknown structural changes using a panel unit root test, which also allows for cross-section dependence. Using these types of tests has two advantages. First, structural breaks in variables can mask the existence of time-varying patterns. Second, the analysis of possible cross-section dependence is especially relevant for highly integrated countries where financial contagion is expected. Moreover, unit root tests can capture the potential existence of highly persistent unobserved drivers for the measured data.

Bai and Carrion-i Silvestre (2009) propose a set of panel unit root statistics that pool the modified Sargan-Bhargava (hereafter MSB) tests[27] for individual series, both taking into account the possible existence of multiple structural breaks[28] and modeling cross-section dependence as a common factor model.[29] The common factors may be nonstationary processes, stationary processes, or a combination of both. The number of common factors is estimated using the panel BIC, as in Bai and Ng (2002).

We have applied an adapted GAUSS code allowing for three breaks, as determined through the Bai and Perron (1998) procedure.[30] When applying this test for our 9-country panel, we also find strong evidence of multiple structural breaks affecting most of the variables analyzed, differing in number and position for each country, as shown in Table 1. The procedure detects at least one structural break in each time series. This evidence suggests that empirical analyses neglecting the presence of structural breaks may omit relevant information.

Table 2 presents the panel-based unit root test results. When allowing for common factors and structural breaks, the results favor the non-stationarity of the variables, as the null hypothesis of a unit root cannot be rejected in most cases, no matter the model or the specification used.

According to Bohn (2007), a stochastic debt is consistent with its corresponding IBC if the series is stationary at any finite number of differences. In our context, this proposition indicates that if any finite difference of *nfar* is stationary, the *nfar* positions are consistent with solvency. As pointed out by Bohn (2007), the intuition is that if *nfar* is mth-order integrated, its n-period-ahead conditional expectation is a polynomial of at most order m. However, the discount factor in the transversality condition grows exponentially with *n*. Therefore, since exponential growth dominates the polynomial growth of any order, *nfar* grows slower than the discount factor in the transversality condition as long as *nfar* is integrated of any finite order.

Table 2. Bai and Carrion-i-Silvestre (2009) panel unit root test with common factors and structural breaks (1973–2021)

Model 2. Trend break model										
	Test statistics			**Simplif. tests statics**						
	Z	P_m	P	Z	P_m	P	T	N	m	fr
$xn_{i,t}$	−0.984	−0.368	15.793	−0.984	−0.368	15.793	49	9	3	27
$nfa_{i,t}$	−1.341	0.986	23.915	−0.70	0.874	23.244	49	9	3	27
$nfac_{i,t}$	−1.728**	3.099***	36.597***	−1.801**	3.137***	36.824***	49	9	3	27
$finopen_{i,t}$	−1.715**	1.381	26.286*	−0.951	1.355	26.131*	49	9	3	27
$kaopen_{i,t}$	−1.549	1.207	25.244	−1.123	1.170	25.018	49	9	3	27

Notes: Z, P and Pm denote the test statistics proposed by Bai and Carrion-i-Silvestre (2009). Z and P_m follow a standard normal distribution and their 1%, 5% and 10% critical values are, −2.326, −1.645, −1.282, and 2.326, 1.645 1.282, respectively. P follows a Chi-squared distribution with 2N degrees of freedom with critical values *34.805, 28.869 and 25.989*, at 1%, 5% and 10% respectively). The number of common factors are estimated using the panel Bayesian information criterion proposed by Bai and Ng (2002). Z^*, P^* and P_m^* refer to the corresponding statistics obtained using the p-values of the simplified MSB statistics. The null hypothesis of a unit root is rejected at $^*p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$ significance level, respectively, if the statistic is greater than the upper level.

Table 3. Variable $baas_t$. Kim and Perron (2009) and Carrion-i-Silvestre et al., (2009) GLS unit roots tests with multiple structural breaks

	Test statistics	5% critical values
P_T^{GLS}	13.766**	7.675
MP_T^{GLS}	14.026**	7.675
ADF	−4.0760**	−4.043
Z_α	−24.607**	−32.846
MZ_α^{GLS}	−18.007**	−32.846
MSB^{GLS}	0.165**	0.122
MZ_t^{GLS}	−2.978**	−4.043
Break dates	1974 / 1979 /2007	

Notes: * denotes significance at $p < 0.10$, $^{**}p < 0.05$, and $^{***}p < 0.01$ significance level, respectively. The critical values were obtained from simulations using 1,000 steps to approximate the Wiener process and 10,000 replications. Note that for the MSB and MP_T tests the null hypothesis of a unit root is rejected in favor of stationarity when the estimated value of the statistic is smaller than the critical value.

Finally, for the variable $baas_t$, which has no panel dimension, we have applied the univariate GLS-based unit root test statistics proposed in Kim and Perron (2009) and extended by Carrion-i Silvestre et al. (2009). The advantage of this test is that it solves problems that previous standard unit root tests with structural breaks had, which assumed that the break occurs only under the alternative hypothesis of stationarity. This test allows for multiple breaks under null and alternative hypotheses at an unknown time. In this case, the null hypothesis of a unit root (with structural breaks) can be rejected at the 5% significance level as shown in Table 3. Accordingly, we can conclude that the variable is stationary.

Thus far, the results of the panel unit root test point to the non-stationarity of nfa_{it} (stock variable) while the nx_{it} ratio is stationary (flow variable). Different orders of integration in intertemporal external budgetary constraint variables can be tackled through multicointegration techniques (Camarero et al. 2013). The interplay between stationary and nonstationary variables, as well as potential structural changes, is a non-trivial feature. A stationary dependent variable, without further restrictions, could be most likely driven by a nonstationary state, such as one following a random walk.[31] Moreover, structural breaks could indicate a time-varying relationship in the model. Finally, the panel evidence can be misleading as it may hide significant country heterogeneity in the dynamic adjustment undertaken by the different euro area countries. Therefore, a proper analysis calls for a more refined specification in a more flexible reaction function. In our case, we do not assume that the transition equation follows a random walk but estimate it as a component of the hyperparameter vector.

5.3 Panel TVP model estimation of the reaction function

In this section, we estimate the empirical state-space specification described in Section 4. According to Hamilton (1994a) among others, the state-space representation of a model provides a flexible framework for modeling time-varying parameters. More specifically, the estimation of the model has been performed using *sspaneltvp*, a complete toolbox developed by the authors in GAUSS to estimate, using the Kalman filter algorithm, a fully fledged state-space time-varying parameter framework for panel time series.

Academic interest in time-varying parameter models for panel time series has expanded rapidly during recent years, providing interesting econometric approaches. Chen (2019) suggests estimating latent group structures in heterogeneous time-varying coefficient panel data models, avoiding poor estimation accuracy in heterogeneous panels when the time series length is short. Feng et al. (2021) use B-splines to model time-varying coefficients for panel data with unobservable multiple interactive fixed effects correlated with the regressors. Alternatively, Custodio João et al. (2023) propose a dynamic clustering model also in a framework of a multivariate panel data model. More recently, Zhao et al. (2025) also address the issue of time-varying coefficients in panel data considering both fixed effects and spatial autoregressive errors.

Our proposed state-space framework provides a fully-fledged approach that can be easily estimated using the Kalman filter (Kalman, 1960), which permits the estimation of the underlying states, in our case, the time-varying parameters, and also the hyperparameters driving their evolution over time.

We contribute to previous work in three directions: first, we develop a seminal GAUSS code written by Hamilton to fit a time-varying multiparameter model; second, we extend the Kalman filter estimation of a single-country model to panel data; third, we adapt the transition equation to include control inputs. Both fixed parameters in the measurement equation and control variables can be common or idiosyncratic (country-specific) parameters.

Essentially, *sspaneltvp* extends the basic model as shown in Hamilton (1994a) into a panel time-varying parameter model, allowing for both fixed (either common or country-specific) and varying components. Under specific conditions, this setting becomes a mean-reverting model, where the fixed mean parameter may include a deterministic trend. Regarding the transition equation, we allow for estimating different autoregressive alternatives and control instruments whose coefficients can be set up either common or idiosyncratic (this is particularly interesting for detecting heterogeneity among individuals, i.e., countries, to common shocks). Furthermore, the GAUSS code allows for restrictions to the variances of both the transition and measurement equations. Other applications using earlier versions of this code can be found in Camarero et al. (2021), Camarero et al. (2020), Paniagua et al. (2017a) and Paniagua et al. (2017b)

The empirical specification was chosen using likelihood-ratio tests among competing alternatives nested in the general model. Often, the likelihood-ratio test statistic is expressed as a difference between the log-likelihoods in the special case that the null hypothesis is true and the alternative (unrestricted) case:

$$\lambda_{LR} = -2 \cdot \left[\ln \left[\sup_{\theta \in \Theta} \mathcal{L}(\theta) \right] - \ln \left[\sup_{\theta \in \Theta} \mathcal{L}(\theta) \right] \right] \tag{23}$$

Multiplying by -2 ensures mathematically that (by Wilks' theorem) $\lambda_{L.R.}$ is asymptotically χ^2-distributed if the null hypothesis is true. The null hypothesis H_0 assumes that the nested

Table 4. Likelihood ratio test results for model specification

Model	Restrictions	Parameters	Hypothesis	LogLikel	LR	df	p-value	Result
1	Full model	99						
2	$\phi_i = \phi$	91						
3	$\phi_i = \phi$ $sigv_i = sigv$	83		−445.67186				
4	$\phi_i = \phi$ $sigv_i = sigv$ $\mu_{2i} = \mu_2$	75	H_0: Model 4 H_1: Model 3	−446.12673	0.90974	8	0.9988***	Accept
5	$\phi_i = \phi$ $sigv_i = sigv$ $\mu_{2i} = \mu_2$ $\gamma_{kc,i} = \gamma_{kc}$	59	H_0: Model 5 H_1: Model 4	−461.00760	29.76174	16	0.01928	Reject
6	$\phi_i = \phi$ $sigv_i = sigv$ $\mu_{2i} = \mu_2$ $\gamma_{kc,i} = \gamma_{kc}$ $\beta_{k,i} = \beta_k$	27	H_0: Model 6 H_1: Model 5	−519.59899	117.18278	32	0.0000	Reject

Notes: p-values. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

(restricted) and full (unrestricted) models fit the data. If we accept H_0, we should use the nested (restricted) model. Under the alternative (H_A), the unrestricted model fits the data significantly better than the nested model. Then, we can reject the null hypothesis and conclude that the full model offers a significantly better fit and that the imposed restriction would not be adequate.

In Table 4, we present the results of the sequential application of the log-likelihood ratio test, starting with the unrestricted model. Model 4 is the chosen specification that includes an idiosyncratic impact of global risk aversion (interacted by financial openness) and a common parameter for our *de iure* measure of financial openness. The autoregressive parameter and the variance of the transition are also common for all elements in the space vector. In the case of the measurement equation, all mean fixed parameters are country-specific.

The results for the maximum likelihood estimation of the hyperparameter vector are reported in Table 5. According to our model, external sustainability occurs when there is a dynamic relationship between net exports and a country's net asset position. As we have already stated, sustainability implies a negative relationship between the two variables due to how the stock variable is defined. In the empirical model, we allow for estimating both a fixed mean (idiosyncratic) coefficient and its time-varying counterpart.

Concerning the measurement equation, the intercept is significant (and negative) for all the peripheral countries analyzed except for Ireland, the only peripheral country exhibiting surpluses on the verge of the G.R. The intercept is positive and significant only for Germany, among the core EMU countries. The autoregressive coefficient is also highly significant but heterogeneous, as its value ranges between 0.332 (Greece) and 0.903 (Germany). The effect of the variable nfa_{t-1} on net exports nx_t as captured by the fixed parameter β_{2i} is different for each country and significant for only three. It is negative in the cases of Austria and Portugal but positive in Greece, implying an unsustainable path in its external position.[32] However, this initial outcome needs to be qualified. As European financial integration has increased access to funding, it has also strengthened the ability of EU countries to incur external imbalances over time. To account for this effect, we define a new variable, $nfac_{t-1}$, which weights nfa_{t-1} by a measure of financial openness, which we use as an additional regressor. In this case, its coefficient (parameter β_{3i} in equation (20)) is smaller but negative and significant in the cases of Germany, Greece, and Spain; for Austria and Portugal, the sign is positive. The cyclical variables $yvar_t$ and $mvar_t$ have idiosyncratic parameters in the panel that are both significant in most cases and exhibit the expected signs.

Table 5. State-space external reaction function 1973-2021. Selected EMU countries

	AUT	BEL	FRA	DEU	GRC	IRE	ITA	POR	ESP
Measurement Equation									
$xn_{i,t} = \beta_{0i} + \beta_{1i} xn_{i,t-1} + \left(\beta_{2i} + \beta_{3i} ln(finopen_{i,t-1})^{-1} + \xi_{i,t}\right) nfa_{i,t-1} + \gamma_{1i} yvar_{i,t-1} + \gamma_{2i} mvar_{i,t-1} + \omega_{i,t}$									
β_{0i}	-0.054	0.126	0.025	0.872^{***}	-5.060^{***}	2.746^{*}	-0.661^{**}	-3.955^{***}	-2.244^{***}
	(0.311)	*(0.182)*	*(0.151)*	*(0.255)*	*(1.086)*	*(1.609)*	*(0.308)*	*(0.889)*	*(0.484)*
β_{1i}	0.808^{***}	0.840^{***}	0.718^{***}	0.903^{***}	0.332^{***}	0.728^{***}	0.830^{***}	0.648^{***}	0.764^{***}
	(0.111)	*(0.067)*	*(0.097)*	*(0.053)*	*(0.122)*	*(0.107)*	*(0.081)*	*(0.081)*	*(0.082)*
β_{2i}	-1.085^{**}	-0.114	0.116	0.141	0.498^{**}	-0.820	0.294	-0.201^{**}	0.129
	(0.549)	*(0.173)*	*(0.380)*	*(0.264)*	*(0.187)*	*(0.594)*	*(0.409)*	*(0.104)*	*(0.143)*
β_{3i}	0.043^{*}	0.003	-0.008	-0.014^{*}	-0.020^{**}	0.044^{*}	-0.020	0.007	-0.013^{**}
	(0.023)	*(0.008)*	*(0.016)*	*(0.009)*	*(0.008)*	*(0.029)*	*(0.017)*	*(0.005)*	*(0.006)*
$\gamma_1 i$	-0.348	-0.027^{***}	-1.772	-3.365^{**}	-0.077	-6.623^{**}	3.948^{*}	-1.928	0.787
	(2.674)	*(0.0)*	*(1.916)*	*(1.599)*	*(2.228)*	*(3.829)*	*(2.223)*	*(1.968)*	*(1.219)*
$\gamma_2 i$	-0.469	-1.325^{***}	-0.824^{**}	-0.955^{**}	-0.049	-6.398^{***}	0.158	-1.720^{***}	-0.805^{**}
	(0.672)	*(0.297)*	*(0.441)*	*(0.419)*	*(1.068)*	*(1.878)*	*(0.555)*	*(0.605)*	*(0.371)*
sigw	0.881^{***}	0.732^{***}	0.730^{***}	0.716^{***}	1.092^{***}	5.361^{***}	1.005^{***}	1.438^{***}	0.965^{***}
	(0.089)	*(0.074)*	*(0.074)*	*(0.072)*	*(0.111)*	*(0.542)*	*(0.102)*	*(0.146)*	*(0.098)*
State Transition equation									
$\xi_{i,t+1} = \phi \xi_{i,t} + \mu_{1i} [baas_t * kaopen_{i,t}] + \mu_2 kaopen_{i,t} + \upsilon_{i,t+1}$									
ϕ	0.894^{***}								
	(0.013)								
μ_{1i}	0.006	-0.005^{**}	-0.008^{*}	-0.006	-0.017^{***}	0.002	-0.008^{*}	-0.006^{***}	-0.006^{***}
	(0.009)	*(0.002)*	*(0.005)*	*(0.006)*	*(0.003)*	*(0.011)*	*(0.005)*	*(0.002)*	*(0.002)*
μ_2	0.022^{***}								
	(0.006)								
sigv	0.000								
	(0.001)								
Number of observations: 49									

Notes: standard deviations in parentheses. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

As for the transition equation, the degree of persistence is high (measured by a coefficient ϕ of 0.894) but smaller than one. In addition, the parameter μ_{1i} suggests strong asymmetries in the reaction to global risk aversion, measured by the interaction of global risk aversion in BAA spread and the *de iure* composite for countries' financial openness or $baas_t * kaopen_{i,t}$, that have a significant impact on peripheral countries except Ireland compared to core EMU members. Finally, the parameter of financial openness μ_2, common for the panel, points to a positive influence of international financial integration in adjusting external debt. Compared with the rest, Austria and Germany have different behavior, as μ_{1i} is non-significant in these two cases. A plausible explanation is that openness is relevant in both of them. Still, they are perceived as non-risky, so the first component, the interaction between risk and openness, is not significant. For the remaining core countries (Belgium and France) μ_{it} is significant and negative.

The interpretation of the results and, in particular, the role played by the *nfar* position is straightforward: although lagged external debt accumulation triggers the need for an external adjustment, financial openness relaxes the external constraint of the countries. The joint effect is represented by the sum of the two constant parameters plus the time-varying term $(\xi_{i,t})$ in $\left(\beta_{2i} + \beta_{3i} ln(finopen_{i,t-1})^{-1} + \xi_{i,t}\right) nfa_{i,t-1}$ that can be analyzed by looking at equations

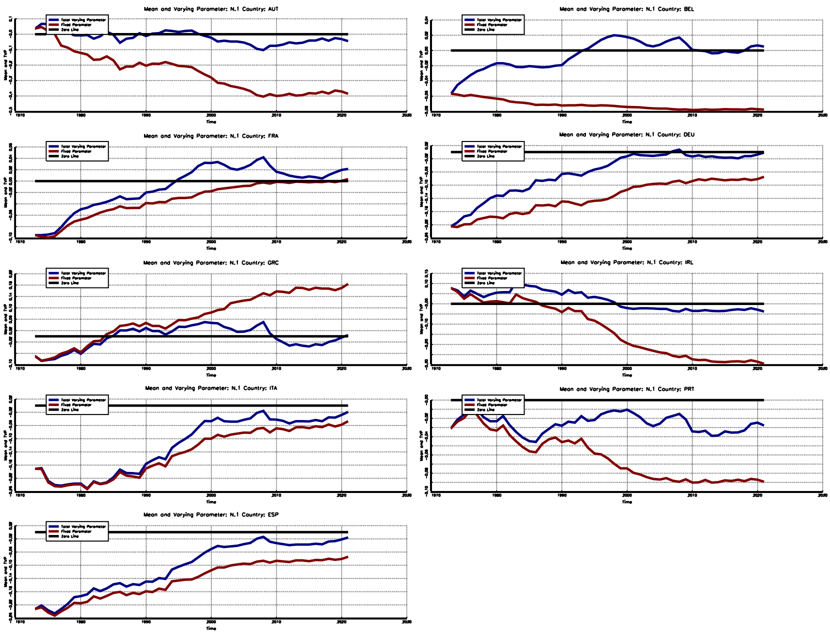

Figure 6. Time-varying external reaction function. Selected EMU countries. 1973–2021.

(21) and (22). This result can be better understood through the visual inspection of Figure 6, where the blue line represents the total effect (constant and time-varying) of the net foreign assets on net exports. This blue line shows the evolution of the time-varying external reaction function for the nine countries. They all have in common that, regardless of the situation at the beginning of the sample, they have converged towards a small negative value at the end of the period. This implies that a process of external adjustment has taken place. As external solvency would imply relatively small (in absolute terms) negative values of the joint reaction, we can discuss each country separately, as their circumstances are quite different. Before analyzing the country heterogeneity using the idiosyncratic reaction functions, it is worth noting that the difference between the red and the blue lines in all the graphs contained in Figure 6 is that the former does not include the time-varying term $\xi_{i,t}$, that is, the effect of the transition equation (22).

We single out two groups of countries: core EU members (Austria, Belgium, France, and Germany) and five peripheral ones (Greece, Ireland, Italy, Portugal, and Spain). Concerning the first group, France has a different behavior than the rest, as the red and blue lines evolve jointly during the whole period, even during the various crises. However, the relevance of the time-varying component is clearly shown in the graphs: in Austria, Belgium, and Germany, the adjustment towards zero is brought by the transition equation. Concerning the role of net foreign assets and their interaction with openness, they are significant at 10% and positive in the cases of Ireland and Austria; negative for Germany (at 10% significance), Greece and Spain, in both at 5%. In the case of output, the cyclical import component, captured by γ_{1i}, is quite heterogeneous, significant, and negative for Belgium, Germany, and Ireland and positive for Italy. In the case of γ_{2i}, is negative when significant and nonsignificant for Austria, Greece, and Italy.

In all cases, except Greece, the fixed component of the reaction function (red line) has a negative value. However, when we add to this fixed component the impact of more significant financial openness in some countries, such as Germany or Spain, this component's absolute value decreases, implying less need for reaction. On the contrary, in Austria or Ireland, greater financial openness seems to lead to an adjustment in the trade balance. Greece and Ireland's trajectories go in opposite directions: while Greece gradually moves away from a reaction towards sustainability, Ireland very soon enters negative and increasing values.

The changing component of the reaction parameter (difference between the red and the blue line), captured by the unobservable element in the transition equation, should be added to obtain the total changing reaction parameter. In particular, for Greece and Portugal, especially since 2007, the combination of financial openness and uncertainty requires a trade balance adjustment when facing difficulties in financing external debt. Austria has adjusted since 2000, a small country that has benefited from financial globalization, but France, since 2007, has adjusted, except for the last two years. All countries have generally been pushed towards adjustment by external financial conditions. More indebted countries have been forced to change more strongly. For surplus countries, such as Germany, there is no such pressure. Belgium and France were the only core countries that needed to react strongly after the 2007 crisis. Due to the increase in the $baas_t$ spread, the blue line shows the intensity of this reaction.

For peripheral countries, only in the cases of Greece, Belgium, and France was the time-varying parameter positive before 2010, so the non-varying part of the model (red line) grew systematically above zero. Indeed, Greece only corrected the external imbalances after the country was bailed out during the financial crisis. Note that $kaopen_{i,t-1}$ shows a different sign for Greece and Portugal. While in the latter, a higher financial openness helps to relax the external constraint, in the case of Greece, financial openness worsens the external sustainability, pointing to a capital flight crisis. Portugal has a relatively stable reaction function, but the three fixed parameters are significant. The reaction function was stable (negative) during the whole period, but the size of the reaction decreased with international financial integration and participation in the euro. The time-varying evolution of the external adjustment in response to external debt accumulation is driven entirely by external and non-permanent factors, such as the BAA spread measure of global risk aversion. At least for Portugal, global risk aversion significantly impacts the reaction by tightening it. Finally, regarding the control instruments employed, financial openness (measured by the Chinn-Ito Index) seems to relax, for the Portuguese case, the need to adjust its external imbalance. For Ireland, the heavy entries of foreign capital (represented by positive fixed parameters β_0 and β_1, and non-significant β_2) explain the flat profile of the external reaction function.

Finally, the cases of Italy and Spain display similar behavior. Both start from large negative values of the reaction function, as during the eighties and the beginning of the nineties, the adjustment of external imbalances (in a context of capital controls and low access to the international capital markets) relied on exchange rate devaluations and net imports. The cost of external sustainability was reduced with the Single European Market in 1992 and, therefore, the removal of capital controls. The euro's inception in 1999 completed the international financial integration of the two countries. The monetary union membership seems to have lessened the external restriction for these two countries so that the adjustment size necessary when external debt accumulates has decreased, reaching values between 3–5% at the end of the sample.

6. Conclusions

This paper uses a panel-TVP state-space framework to examine the dynamic external reaction of the net exports to external liabilities accumulation. Using the Kalman filter (KF), we estimate the state-space model for nine euro area countries, namely, Austria, Belgium, France, Germany, Greece, Ireland, Italy, Portugal and Spain. We contribute to previous work in four directions. First, we modify Hamilton's GAUSS code to fit a time-varying multi-parameter model; second, we extend the KF estimation from a single-country model to a panel setting; third, we adapt the transition equation to include control inputs; finally, both fixed parameters in the measurement equation and the control variables are allowed to be either common or idiosyncratic.

Using this empirical setting, we transpose Bohn's fiscal reaction approach to analyze the external sector constraint. To our knowledge, previous empirical attempts to capture this external sector reaction function using other empirical strategies have obtained disappointing outcomes. Yet, our results show a precise interplay between real and financial variables, with the latter increasingly important. Unit root tests show evidence of a different order of integration of the variables involved in the IEBC, pointing to the error correction model approach suggested by Bohn as the golden rule to capture the dynamics between net exports and the net foreign asset position (*nfa*). The net exports ratio is found to be stationary during the period 1970–2015. The other variables involved are generally non-stationary, except for the BAA spread.

The time-varying estimate of the external reaction function helps us to disentangle the variations in the degree of responsiveness over time. Our empirical approach allows us to break down the joint effect into the sum of two constant parameters plus a time-varying term. The five EU countries analyzed were found to adjust their external balance positively to rising stock levels of external liabilities. All countries have in common that, regardless of their initial situation, they all converged towards a small negative value at the end of the period, meaning a process of external adjustment. Moreover, the heterogeneity of the results at the country level highlights the advantages of estimating country-specific external reaction functions when possible. Most coefficients evolve around a fixed parameter with some degree of persistence. The filtered components of imports and income play a common role for all the countries. In addition, one of the most interesting findings of the paper is that external debt accumulation has to be put in relation to the degree of financial openness, which has relaxed in recent decades the external constraint of the euro economies, easing the financing of a country's imbalances through external financial markets. However, financial openness facilitates current account reversals beyond a threshold, triggering a reaction to reduce the external imbalance. Our results show that Greece only corrected its external imbalance after being bailed out during the financial crisis. While in the rest of the peripheral countries, more financial openness helps to relax the external constraint (which is captured both by the measurement and the state equations), in the case of Greece, financial openness worsens its external sustainability, pointing to a capital flight crisis. Finally, changes in the risk appetite of the markets also play a key role at specific points in understanding the dynamics of the external adjustment.

While our paper presents a state-space framework that can be efficiently estimated using the Kalman filter, several important methodological challenges in heterogeneous panels merit further investigation. The rapidly expanding literature on time-varying models in panel data offers promising avenues for future research. For instance, estimating latent group structures in time-varying-coefficient models suggests potential extensions for identifying unobserved heterogeneity. Other important considerations include the incorporation of interactive fixed effects and the treatment of cross-sectional autoregressions. Additional methodological challenges worth exploring include appropriate clustering techniques for heterogeneous panels, refined approaches to spatial correlation, and specialized estimators when dealing with time-trending regressors. These extensions would further enhance the flexibility and applicability of time-varying parameter models in panel data analysis.

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Notes

1 Schoder et al. (2013) find empirical evidence suggesting that introducing the euro is associated with a regime shift from sustainability to the unsustainability of external debt accumulation. Since the launching of the euro, we have witnessed periods of pronounced widening of external imbalances between peripheral and core countries of the euro area. Blanchard and Milesi-Ferretti (2010), or Jaumotte and Sodsriwiboon (2010) stress that the significant rise of external borrowing in some EMU countries would have reflected bubble-driven booms caused by overly optimistic growth prospects.
2 See https://econweb.ucsd.edu/ jhamilto/software.htm for the original code.
3 See Camarero et al. (2025).
4 We use the different terms indistinctly in this text.
5 See Lane and Milesi-Ferretti (2002) and Gourinchas and Rey (2007) for a complete discussion of the longer-term relationship between the US net export deficit and revaluations of the US net foreign asset position.
6 In this paper, we consider the notion of "external sustainability" as synonymous with "external solvency."
7 An external crisis can take several forms, such as a (partial) default on external liabilities, debt restructuring, or recourse to international financial assistance.
8 In textbook examples, there is no distinction between i_A and i_L because they assume only one traded asset. However, in the real world, there are sometimes significant differences (Gourinchas and Rey, 2007).
9 This assumption can be considered mild to some extent, as we are dealing with countries within a monetary union. However, the valuation effects on other monetary areas can be significant.
10 In this setup, it is possible to distinguish different degrees of sustainability, i.e., if $\beta_{1,0} = 1$ we have strong sustainability whereas weak sustainability is met when $\beta_{1,0} < 1$ (Quintos, 1995).
11 Note that although Bohn's assertions are referred initially to public finance, they can be easily applied to external imbalances.
12 See Trehan and Walsh (1991), who found evidence supporting the stationarity of the US external investment position. Other examples include Wickens and Uctum (1993), Ahmed and Rogers (1995), and Liu and Tanner (1996).
13 This approach has been adopted by several authors, including Lane and Milesi-Ferretti (2007), Gourinchas and Rey (2007), and IMF (2005).
14 Notable examples include Husted (1992), Fountas and Wu (1999), Irandoust and Sjöö (2000), Irandoust and Ericsson (2004), Arize (2002), Narayan and Narayan (2005), Herzer and Nowak-Lehmann (2006), Hamori (2009), and del Barrio-Castro et al. (2015, 2019).
15 The idea originally proposed in Bohn (1998).
16 Even if this concept does not exist "stricto sensu" in the external accounts, we define it here as the current account balance excluding capital income.
17 See footnote 2 in Bohn (2007).
18 Note that we are using the term primary balance as transposition of the terminology coined by Bohn to establish an analogy between the fiscal reaction function and a external reaction function. In our case, we assimilate this term to the trade balance
19 A more-in-depth description of this framework can be found in Camarero et al. (2020).
20 See, for instance, Hodrick and Prescott (1981).
21 Note that an increasing external debt would mean a negative position regarding the stock of net foreign assets.
22 This specification including the lagged net exports on the right-hand side was preferred based on the strong evidence of persistence of the dependent variable in all alternative specifications and trying different estimators.
23 Jahan and Wang (2016) provide an alternative measure of financial openness that includes a breakdown of openness for various subcategories of capital flows. However, their series started in 1996.
24 As argued by Bajo-Rubio et al. (2014), restricting the span of our data from 2008 onward would tend, in a fixed parameter model, to bias the results, given that external imbalances have been significantly reduced since the onset of the crisis.
25 We have downloaded it from http://web.pdx.edu/ ito/Chinn-Ito_website.htm.
26 A popular alternative to account for global risk aversion is the CBOE Volatility Index, which measures the stock market's expectation of volatility implied by S&P 500 index options, calculated and published by the Chicago Board Options Exchange. However, this variable was not available for the whole period. Moreover, while BAA spread measures risk appetite, a variation

in the implied volatility in a market may stem from a change in the quantity of risk in this market and not necessarily from a change in the investor's risk aversion.

27 See Sargan and Bhargava (1983).

28 Adapting Bai and Perron (2003) methodology to a panel data framework.

29 Following Bai and Ng (2004).

30 See Bai and Carrion-i Silvestre (2009) for details.

31 An essential part of the literature on time-varying parameter models would make this choice.

32 Similar results were obtained by Engel and Rogers (2006) for the U.S. economy.

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