STRUCTURAL AND STOCHASTIC ASYMMETRIES WITHIN FIXED EXCHANGE RATES SYSTEMS

Jean-Sébastien Pentecôte

Associate Professor, CREM–UMR CNRS 6211

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Abstract:

Bayoumi and Eichengreen's (BE, 1994) article remains influent in the empirics of the core-periphery view of fixed exchange rate agreements. They rely on the basic AS-AD framework in order to identify supply and demand shocks through long-run restrictions in vector autoregressions. It is shown here how the previous authors depart from Blanchard and Quah's (BQ, 1989) factorization. Contrary to BE's premise, relaxing the assumption of shocks of equal size is not just a matter of scale. The properties of a fixed exchange regime can be severely distorted, especially as regards the correlation between shocks and their relative size. Furthermore, zero-constraints on either instantaneous or long-run impulse responses provide identical results given the peculiar specification of the VAR process usually considered in the literature. Finally, alternative methods are compared to derive the slope coefficients of the underlying AS and AD curves. It appears that only non-zero restrictions on VARs imply values consistent with those derived from New-Keynesian models. Monthly data for the founder countries of the euro area over 1996-2008 illustrate all these points.

Keywords: fixed exchange rates, core-periphery, long-run restrictions, structural VARs

JEL codes: C32, E13, F33

Affiliation: CREM–UMR CNRS 6211, Faculty of Economics, University of Rennes 1, 7 place Hoche, CS 86514, 35065 Rennes Cedex, France. Tel: (33) 2 23 23 35 56. Fax: (33) 2 23 23 35 99. Email: jpenteco@univ-rennes1.fr

1. Introduction

Disentangling the empirical properties of the macroeconomic shocks hitting a set of countries is a crucial issue in exchange rate economics. The stochastic dependencies exhibited among countries influence their choice to peg their currency to a foreign anchor or even to join a monetary union. They help explain regional exchange rate agreement like in the European Union (Bayoumi and Eichengreen (1992)) as well as the polarization of the international monetary system around a few currencies (Bayoumi and Taylor (1995)).

According to the literature on optimum currency areas, sharing the same currency and committing to a common monetary policy critically depends on the nature and the size of the macroeconomic disturbances when there are no substitutes for exchange rate adjustments. Fixed exchange rates should be preferred when common (symmetric) shocks dominate the idiosyncratic ones and/or they call for symmetric responses.

In a series of empirical works, Bayoumi and Eichengreen (1992, 1994) popularized the core-periphery view of the functioning of hard pegs. They rely on the textbook Aggregate Supply-Aggregate Demand (AS-AD) model to show how nation-wide supply and demand shocks can be extracted from the joint autoregressive output growth-inflation dynamics. The former is described by a finite-order bivariate VAR process. The underlying structural shocks can then be recovered from the VAR residuals according to a given set of identifying restrictions. Bayoumi and Eichengreen (BE later on) refer to the procedure developed by Blanchard and Quah (1989, BQ elsewhere) since it is assumed the long-run neutrality of real output to demand shocks.

In the recent years, the pursuit of the monetary unification process in an enlarged European Union has revived the debate around the asymmetric functioning of the euro area itself.

Almost all the empirical studies on this issue refer to VAR identification based on long-run restrictions (Fidrmuc and Korhonen (2006) for a survey). They aim at assessing the nature and the extent of stochastic asymmetries among a set of countries which (ambition to) share the same currency and the foregoing monetary policy. Asymmetry is usually gauged through correlation between each country pair at two levels: the nature of macroeconomic shocks and the adjustment process to those disturbances.

However, the decomposition of the shocks of the VAR they consider departs from the more familiar BQ approach in one major respect. BE indeed relax the assumption of equal (unitary) variances of the structural innovations. They are interested not only in the correlation between domestic and foreign disturbances, but also in the relative size of demand and supply shocks in each country. In their view, departing from the usual normalization assumption is essential because the size of shocks matters for assessing the extent of asymmetries within a monetary union.

They further advocate that decomposing the correlation matrix of the VAR residuals rather than the variance-covariance matrix itself is inconsequential to the measure of correlation coefficients. They state that: "*These two normalization gave almost identical paths for the shocks, except for a scaling factor, and hence are used interchangeably*" (BE (1994), p. 816).

This may be one of the reasons why the BE procedure has not been strictly followed in the literature devoted to the Eastern enlargement of the euro area. On one hand almost all these studies refer to the standard textbook AS-AD model like BE (1992) as a way to justify the long-run restriction put on the (absence of) response of output to shocks from the demand side. On the other hand, the same studies use a set of identification constraints similar to BQ in order to obtain the structural form of the VAR process.

The aim of this study is thus to give a critical appraisal of the relevance of long-run restrictions in structural VAR models within the textbook AS-AD theoretical framework.

This question is often viewed as one of the masterpiece of the identification problem in structural VAR modeling. It is only recently that new answers have been proposed to this broader issue (Rubio-Ramirez, Waggoner, and Zha (2010)). However, severe doubts remain about the usefulness of non-linear constraints on the VAR parameters arising from the long-run properties of the most popular macroeconomic models. This clearly involves the decomposition between permanent and transitory disturbances.

The discussion proceeds as follows. Section 2 questions the “equivalence principle” suggested by BE when the decomposition of structural shocks is based on the *correlation* matrix of the VAR residuals. It is shown that the path followed by *each* of the structural shocks will be unchanged only if an orthonormal matrix is chosen to ensure the transition between the BQ and the BE factorizations. There is no reason to believe, contrary to BE’s conjecture, that the latter is always the identity matrix. Instead, it is established that such transition matrix is defined only up to some appropriate rotation. It appears, Section 3 shows how the auxiliary equations for the VAR identification can directly be derived from the slopes of the AS and the AD curves in long- and/or the short-term. A special emphasis is put here on the equivalence of the BQ approach with competing zero- and sign-restrictions on the impact response to either permanent or transitory impulses, consistently with the AS-AD diagram. Section 4 gives an empirical illustration of our results. We study asymmetries among the European countries (including Greece). Monthly HCPI and IPI data cover the 1996:01-2008:12 period. Our results confirm that switching from the BQ to the “unadjusted” BE decomposition may have severe consequences about the relative size and the correlation of shocks depending, thereby modifying the core-periphery view of the euro area.

Furthermore, our empirical findings confirm that performing the BQ factorization yields exactly the same results as a Choleski decomposition given the particular VAR setting inherited from BE. To this view, resorting to some long-run neutrality assumption would add little, if any, to the identification problem of the underlying AS-AD theoretical model as well as the empirical issue of shock asymmetry under a common fixed exchange rate agreement. Still, when relying on the AS-AD diagram, our estimates reveal that sign restrictions seems to offer a better alternative to restrictions due to some Wold ordering of the variables. The former constraints lead to slope coefficients of the Lucas-type supply function and the Phillips-type inflation-output growth relationship closer to existing estimates from popular New-Keynesian models. Our agnostic approach of the VAR-process for the output gap and HCPI inflation indeed gives strong support to the view of very flat aggregate supply like aggregate demand curves, though exhibiting significant discrepancy from one euro Member to another. Section 5 concludes.

2. Long-run output neutrality and the size of shocks

2.1. The Blanchard and Quah (1989) principle and the auxiliary assumptions

Let us consider that the reduced form of the price-output dynamics in a given country is given by the following *p*-order bivariate VAR process:

$$\begin{pmatrix} g_t \\ \pi_t \end{pmatrix} = \begin{pmatrix} a_{11}^1 & a_{12}^1 \\ a_{21}^1 & a_{22}^1 \end{pmatrix} \begin{pmatrix} g_{t-1} \\ \pi_{t-1} \end{pmatrix} + \cdots + \begin{pmatrix} a_{11}^p & a_{12}^p \\ a_{21}^p & a_{22}^p \end{pmatrix} \begin{pmatrix} g_{t-p} \\ \pi_{t-p} \end{pmatrix} + \begin{pmatrix} \varepsilon_{1t} \\ \varepsilon_{2t} \end{pmatrix}, \qquad (1)$$

which can be written as:

$$\mathbf{z}_t = \sum_{i=1}^{p} \mathbf{A}_i L^i \mathbf{z}_t + \boldsymbol{\varepsilon}_t \tag{2}$$

$\mathbf{z}_t$ *is the vector of* the (first-order log-difference) of the economic activity index (g_t) and the (first-order log-difference) of the price index (π_t) at the date t. The terms a_i^0, a_{ij}^k are the parameters of interest of the model. L stands for the lag operator. The vector of VAR residuals $\boldsymbol{\varepsilon}_i(t)$ follows a white noise process with covariance matrix: $\Sigma = \begin{pmatrix} \sigma_1^2 & \sigma_{12} \\ \sigma_{21} & \sigma_2^2 \end{pmatrix}$. All deterministic terms have been removed.

Provided that the VAR (p) is invertible, the corresponding VMA (∞) form is given by:

$$\mathbf{z}_t = \left(\mathbf{I} - \mathbf{A}(L)\right)^{-1} \boldsymbol{\varepsilon}_t \,. \tag{3}$$

As a second step, the identification procedure is used to derive the (“structural”) innovations from the residuals after the estimation of the VAR for each “country”: the candidate one and the euro area itself. Four structural shocks are thus isolated according to whether they relate to the supply or to the demand side and whether they are common to the single currency area or specific to the candidate country.

For the applicant country as for the reference area, the VAR residuals are initially expressed as a linear combination of the structural innovations:

$$\boldsymbol{\varepsilon}_t = \mathbf{C}_{BQ} \boldsymbol{\eta}_{BQ,t}, \quad \forall k \in IN. \tag{4}$$

$\mathbf{C}_{BQ}$ is the lower-triangular matrix consistent with the Blanchard-Quah identification assumptions while $\boldsymbol{\eta}_{BQ,t}$ is the corresponding vector of (structural) innovations.

The moving average representation becomes:

$$\begin{pmatrix} g_t \\ \pi_t \end{pmatrix} = \left(\mathbf{I} - \mathbf{A}(L)\right)^{-1} \mathbf{C}_{BQ} \begin{pmatrix} \eta_{BQ,t}^{P} \\ \eta_{BQ,t}^{NP} \end{pmatrix} \tag{5}$$

with $\eta^{P}_{BQ,t}$ the “permanent” (aggregate supply), and $\eta^{T}_{BQ,t}$ the transitory (aggregate demand) shocks, respectively.

BQ’s (1989) identification constraints lead to the following system:

$$\begin{cases} E\left(\boldsymbol{\eta}_t \boldsymbol{\eta}_t{}'\right) = \mathbf{I}_2 \\ \left(\mathbf{I} - \mathbf{A}(1)\right)^{-1}\mathbf{C}_{BQ} = \begin{pmatrix} . & 0 \\ . & . \end{pmatrix} \end{cases} \tag{6}$$

The first equation leads to non-correlated innovations with unit variance. The second condition implies the lack of a permanent (over an infinite horizon) effect of demand shocks on output. Therefore the long-run impact matrix $\left(\mathbf{I} - \mathbf{A}(1)\right)^{-1}\mathbf{C}_{BQ}$ must be lower-triangular.

Normalization of the variance of the innovations is common practice in structural vector autoregressions. However Bergman (2005) shows the shape of the impulse response functions can be very sensitive to the variance ratio of these stochastic components. Simulations on a bivariate VAR similar to (1) indeed reveal a puzzling positive impact of permanent (from the production side according to the author) shocks to price level when demand shocks largely dominate their supply counterpart. By contrast, this response is consistent with the AS-AD textbook model when the variance ratio is constrained to unity. The relative contribution of the structural shocks to the forecast error variance of the endogenous variables vary also substantially depending on whether the structural shocks are assumed to have equal variance or not.

This issue is also central to the empirical assessment of fixed exchange rate regimes. As stressed by BE (1992), what matters is not only the side from which asymmetries dominate, but also the relative size of the so-called supply and demand disturbances. To this end, the procedure of VAR identification should allow one to compute the variance of the structural shocks in a given country. It remains however unclear how the series of shocks and the related correlation coefficients may be influenced.

2.2. *Bayoumi-Eichengreen vs Blanchard-Quah: only a matter of scale?*

Let us assume like BE that the reduced form of the price-output dynamics in a given country is described by the following *p*-order bivariate DSVAR(*p*) process[1] in the matrix form:

$$\mathbf{z}_t = \sum_{i=1}^{p} \mathbf{A}_i \mathbf{L}^i \mathbf{z}_t + \boldsymbol{\varepsilon}_t \tag{7}$$

Here $\mathbf{z}_t = \left((1-L)y_t, (1-L)p_t\right)'$ is the vector of the (first log-difference) of output and the (first log-difference) of the price level in some period *t*. *L* is the lag operator. VAR residuals $\boldsymbol{\varepsilon}_t$ are white noise processes with covariance matrix $\boldsymbol{\Sigma}_{\varepsilon}$.

Provided that the VAR (p) is invertible[2], the corresponding VMA(∞) form is given by:

$$\mathbf{z}_t = \left(\mathbf{I}_2 - \sum_{i=1}^{p} \mathbf{A}_i \mathbf{L}^i\right)^{-1} \boldsymbol{\varepsilon}_t \,. \tag{8}$$

with $\mathbf{I}_2$ the conformable identity matrix. In a second step, an identification procedure is used to derive the ("structural") innovations from the residuals after the estimation of the VAR.

As noticed earlier, recovering the structural shocks from the VAR residuals plays a crucial role if one is interested, like BE, in the correlation between foreign and domestic – either permanent or transitory – shocks to gauge the suitability of a currency area among any country pair.

[1] Dupaigne et al. (2007) discuss on potential bias arising from DSVARs.

[2] The non-invertibility issue is disregarded here (Blanchard and Quah, 1993, for a discussion).

However, BE propose to decompose the correlation matrix Γ_e of VAR residuals instead of the covariance matrix itself. Recall first that:

$$\mathbf{\Sigma}_{\varepsilon} = \mathbf{D\Gamma_{\varepsilon}D}' \tag{9}$$

D is the diagonal matrix of the standard-errors of the VAR residuals.

Decomposing the correlation matrix Γ_e as initially suggested by BE (1992) yields another matrix $\boldsymbol{C}_{BE}$ leading to the following factorization:

$$\mathbf{\Sigma}_{\varepsilon} = \mathbf{D}\left(\mathbf{C}_{BE}\mathbf{C}'_{BE}\right)\mathbf{D}' \tag{10}$$

Another sequence of structural innovations is then given by:

$$\boldsymbol{\varepsilon}_t = \mathbf{C}_{BE}\boldsymbol{\eta}_{BE,t} = \mathbf{C}_{BE}\begin{pmatrix} \eta^{P}_{BE,t} \\ \eta^{T}_{BE,t} \end{pmatrix} \tag{11}$$

Consider however an orthogonal matrix **Q** such that:

$$\mathbf{\Sigma}_{\varepsilon} = \left(\mathbf{DC}_{BE}\mathbf{Q}\right)\left(\mathbf{Q}'\mathbf{C}'_{BE}\mathbf{D}'\right) \tag{12}$$

We can now state our first two main propositions.

Proposition 1: From equations (3) and (9), the BE and BQ factorizations are equivalent (up to a scale factor) if:

$$\mathbf{C}_{BQ} = \mathbf{DC}_{BE}\mathbf{Q} \tag{13}$$

When identity (13) holds, the “BQ” and the “BE” innovations follow the same time path. Unlike BE’s premise, both normalizations cannot be always used “interchangeably” since **Q** has to be chosen conveniently.

Proposition 2: The "transition" matrix **Q** *between the BE and BQ decompositions of permanent and transitory shocks is given by:*

$$\mathbf{Q} \equiv (\mathbf{DC}_{BE})^{-1}\mathbf{C}_{BQ} \tag{14}$$

Having first decomposed the correlation matrix Γ_e *to build* $\mathbf{C}_{BE}$*, one can use (14) to find* **Q**. That *Givens matrix is however not unique as it is defined up to a convenient "rotation". In a bivariate VAR, this leads to eight possible writings for* **Q***.*

Because each of these transformations leads to a new set of "structural" shocks, this should have important implications for empirical studies. Following proposition 2, one must select the particular form of **Q** so that transitory (like permanent) shocks in a given country are undistinguishable under either BQ or BE decomposition schemes. These paths will thus differ only to a scale factor, as stressed by BE.

Proposition 3: Given (10) and (11), the size difference between permanent and transitory shocks in a country depends on **Q***. The "rotated" BE innovations result from a linear combination of the "original" BE shocks. This is crucial for asserting which type of shocks dominates.*

As it will appear from our empirical work below, the choice of the "transition" matrix **Q** matters for the transitory shocks only. To this respect, the ordering of the variables in the VAR process may have severe consequences for the identification of macroeconomic disturbances.

As a byproduct, the variance ratio of the "BE" structural shocks will be modified in case of a noticeable departure of **Q** from the identity matrix. The corresponding "adjusted-BE" disturbances are indeed given by a linear combination of the “original” BE shocks. This may also modify the conclusions about whether the permanent stochastic term dominates its transient counterpart within a given country or not.

It has finally to be mentioned that **Q** itself is not defined uniquely. In what follows, it will be clear that the former can be viewed as a “rotation” matrix such that the orthonormal property is insensitive neither to a transposition operation nor to some appropriate change in the sign of its elements.

In the bivariate case under study, one peculiar matrix has to be picked up among eight possible candidates. Writing $= \begin{pmatrix} a & b \\ c & d \end{pmatrix}$, it has thus to satisfy the following set of conditions:

$$\begin{cases} a = \cos(\theta) \\ b = \pm\sin(\theta) \\ c = \pm b \\ d = \pm a \end{cases}$$

However each of these transformations leads to a new set of “structural” shocks. In our context, this choice is non-neutral to assessing the size and the asymmetry of macroeconomic shocks under a given exchange rate regime.

3. Linking VAR identifying restrictions to an explicit AS-AD model

3.1. Long-run output neutrality with full price indexation

The identification procedure employed to appreciate whether building or enlarging a monetary union is advisable rests on one strong assumption: structural innovations specific to the applicant country and those which monetary union undergoes must be uncorrelated.

Results can then be biased. A similar point has already been discussed by BQ (1989) themselves, followed by other authors such as Wagonner and Zha (2003).

Another major issue lies in aggregation of shocks and time aggregation that could lead to unreliable results from structural VARs because of the correlation between shocks. In order to elucidate the puzzling strong weight of technological shocks in the real business cycle, Cover, Enders, and Hueng (2006) (hereafter CEH) propose a new method of decomposition of the residuals of the VAR. They argue that their procedure has the appealing feature to be consistent with new-classical as well as neo-Keynesian macro-economic models.

Following CEH (2006), a simple version of the AS-AD model is described by the following set of equations:

$$\begin{cases} y_t^S = y_{t|I_{t-1}}^a + \alpha\left(p_t - p_{t|I_{t-1}}^a\right) + u_t^S \\ y_t^D = y_{t|I_{t-1}}^a - \left(p_t - p_{t|I_{t-1}}^a\right) + u_t^D \\ y_t^S = y_t^D = y_t \end{cases} , \tag{15}$$

where y and p, respectively, are the (logarithms of) output and price *levels* of a country and $x_{t|I_{t-1}}^a$ the expected value at time *t* of variable *x* conditional to the information available up to date *t-1*.

Under extrapolative expectations (with a given finite h-ahead period horizon): $x_{t|I_{t-1}}^a = a_x(L) = \sum_{i=1}^{h} a_{x,i} L^i x_t$, the above system translate into the matrix form:

$$\begin{pmatrix} y_t^S \\ y_t^D \end{pmatrix} = \begin{pmatrix} a_y(L) & \alpha\left(1 - a_p(L)\right) \\ a_y(L) & -\left(1 - a_p(L)\right) \end{pmatrix} \begin{pmatrix} y_t \\ p_t \end{pmatrix} + \begin{pmatrix} u_t^S \\ u_t^D \end{pmatrix} . \tag{16}$$

One can already notice the analogy between this system and the (first-order difference) version in (1) where $g_t \equiv y_t - y_{t-1}$ and $\pi_t \equiv p_t - p_{t-1}$. The reduced form from the structural model (16) is:

$$\begin{cases} y_t = \alpha a_y(L) y_t + \dfrac{1}{1+\alpha} u_t^S + \dfrac{\alpha}{1+\alpha} u_t^D \\ p_t = \alpha a_p(L) p_t - \dfrac{1}{1+\alpha} u_t^S + \dfrac{1}{1+\alpha} u_t^D \end{cases} \tag{17}$$

According to CEH, if the system is stable, its VMA(∞) representation can be deduced from (16) and (17):

$$\begin{pmatrix} y_t \\ p_t \end{pmatrix} = \left(\mathbf{I} - \begin{pmatrix} a_{11}(L) & a_{12}(L) \\ a_{21}(L) & a_{22}(L) \end{pmatrix} \right)^{-1} \begin{pmatrix} \dfrac{1}{1+\alpha} & \dfrac{\alpha}{1+\alpha} \\ -\dfrac{1}{1+\alpha} & \dfrac{1}{1+\alpha} \end{pmatrix} \begin{pmatrix} u_t^S \\ u_t^D \end{pmatrix} \tag{18}$$

Although the shock size is normalized, the CEH approach departs from the BQ identification system in two major ways:

- the slope of the aggregate demand (AD) curve is set unity which assumes complete price indexation in the country over an indefinite horizon;
- the assumption that the structural AD shock has no long-run effect on output yields an estimate for α, the slope of the aggregate supply (AS) curve. Indeed, the response of economic activity to the demand shock is given by:

$$\frac{1}{1+\alpha}\left(\alpha\left(1-a_{22}(1)\right)+a_{12}(1)\right)=0 \quad \Leftrightarrow \quad \alpha = -\frac{a_{12}(1)}{1-a_{22}(1)} \tag{19}$$

One striking feature with the CEH identifying equations lies in that it is no longer necessary to impose the domestic shocks to be contemporaneously uncorrelated. So, there is no reason for the corresponding variance-covariance matrix to be diagonal.

This leads to relax one of the most famous “auxiliary” identifying restrictions of the VAR literature. It circumvents one “identification failure” of the structural VAR approach. As underlined by Cooley and Dwyer (1998), such auxiliary assumptions have a dramatic impact on the structural dynamics, although they have no appealing economic interpretation since they do not derive from a well-defined theoretical model. This is typically in line with the misspecification problems raised by *ad hoc* dynamic linear systems like vector auto-regressions (see Cooley and Leroy (1985), and Braun and Mittnik (1993)).

CEH advocate however that fulfilment of the orthogonalization condition is needed if one is interested in the impulse response functions or the forecast error variance decomposition. The second step of their identification procedure amounts to retrieve the underlying BQ innovations by a suitable factorization. The uncorrelatedness assumption can be viewed as an over-identifying restriction.

As concerns the measurement of asymmetry according to exchange rate agreements, this two-step procedure for discovering the structural VAR imply two sequences of identified shocks: the first one being correlated, whereas the second are orthogonal. It is quite ambiguous which of them features the “true” (structural) innovations.

Furthermore, let us assume full price indexation as in (15) and that price and output are first-order difference stationary like in (1). Given our notations, the CEH decomposition of the VAR residuals would imply:

$$\mathbf{A}\varepsilon_t = \mathbf{B}\eta_{BQ,t} \ , \ A = \begin{pmatrix} 1 & -\alpha \\ 1 & 1 \end{pmatrix} \ and \ B = \begin{pmatrix} b_{11} & 0 \\ b_{21} & b_{22} \end{pmatrix}, \tag{20}$$

From (4), we get the following relationship:

$$\mathbf{C}_{BQ} = \mathbf{A}^{-1}\mathbf{B} = \frac{1}{1+\alpha}\begin{pmatrix} b_{11}+\alpha b_{21} & \alpha b_{22} \\ b_{21}-b_{11} & b_{22} \end{pmatrix}. \tag{21}$$

The slope coefficient of the AS curve can be derived directly from the BQ procedure itself as:

$$\alpha = \frac{c_{12,BQ}}{c_{22,BQ}}, \tag{22}$$

where $c_{ij,BQ}$ lies on the *i*-th row and the *j*-th column of matrix C_{BQ}. Results from (19) and (22) should coincide, but the undefined statistical distribution of parameter α precludes from a formal test because of the strong non-linearities.

Comparing the CEH identification strategy with the BQ one raises the question about the appropriate specification of the VAR process. Making use of the AS-AD model as a theoretical background leads authors to put the emphasis on the long-term properties of the dynamic system. However, the short-run dynamics derived from such a framework may also deliver useful information for the identification of the VAR. The usefulness of the long-run restrictions has to be call into question once more.

3.2. Identifying the slopes of the AS-AD curves without long-run restrictions

Criticism against structural VAR analysis with long-run restrictions has a long tradition (Faust and Leeper (1997) among others). The usefulness of zero-constraints on the long-run dynamic multipliers has also been questioned when the system dynamics is described by a Vector Error Correction Model (VECM). Indeed, an obvious generalization of (2) initially suggested by Crowder (1995) is to write:

$$\mathbf{X}_t = \boldsymbol{\mu} + \boldsymbol{\alpha}\boldsymbol{\beta}^T \mathbf{x}_{t-1} + \sum_{i=1}^{p-1} \mathbf{H}_i L^i \mathbf{X}_t + \boldsymbol{\varepsilon}_t, \tag{23}$$

where the (transpose) vectors are defined by $\mathbf{x}_t = (y_t, p_t)^T$, $\mathbf{X}_t = (g_t, \pi_t)^T$, $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$ are the (2X1) matrices (in the bivariate case here) of loading coefficients of the error correction term ($\boldsymbol{\beta}^T \mathbf{x}_{t-1}$) and of (the unique here) cointegrating vector respectively (such that rk$\left(\boldsymbol{\alpha}\boldsymbol{\beta}^T\right)=1$).

When the VECM is invertible, its corresponding VMA(∞) form is:

$$\mathbf{X}_t = \boldsymbol{\rho} + \boldsymbol{\Theta}(L)\boldsymbol{\varepsilon}_t, \tag{24}$$

where $\boldsymbol{\rho} = \boldsymbol{\Theta}(1)\boldsymbol{\mu}$, $\boldsymbol{\Theta}(1) = \boldsymbol{\beta}_\perp \left(\boldsymbol{\alpha}_\perp^T \boldsymbol{\Psi} \boldsymbol{\beta}_\perp\right)^{-1} \boldsymbol{\alpha}_\perp^T$, $\boldsymbol{\alpha}^T \boldsymbol{\alpha}_\perp = 0$, $\boldsymbol{\beta}^T \boldsymbol{\beta}_\perp = 0$, $\boldsymbol{\Psi} = \mathbf{I} - \sum_{i=1}^{p-1} H_i$.

The related structural VMA representation can be written as:

$$\mathbf{X}_t = \boldsymbol{\rho} + \boldsymbol{\Lambda}(L)\mathbf{u}_t, \qquad (25)$$

where $\boldsymbol{\Lambda}(L) = \boldsymbol{\Theta}(L)\mathbf{P}^{-1}$ such that : $\mathbf{P}\mathbf{x}_t = \mathbf{m} + \sum_{i=1}^{p} \mathbf{P}_i L^i \mathbf{x}_t + \mathbf{u}_t,\ E\left(\mathbf{u}_t \mathbf{u}_t^T\right) = \mathbf{I}.$

Following Ribba (1997), if aggregate output is weakly exogenous with respect to the cointegrating vector $\boldsymbol{\beta}$, $\boldsymbol{\Lambda}(0) = \mathbf{P}^{-1}$ is lower triangular, and $\Gamma(1)$ has a second column of zero terms as in the BQ decomposition. This amounts to assume non-Granger causality of output to the price level in the long-run. Under that condition, the vector of loading coefficients is $\boldsymbol{\alpha}_{\perp} = (1,0)$. In other words, the error correction term does not appear in the equation for the output growth g_t. This is also precisely the case in (2).

A similar argument is also raised by Pagan and Pesaran (2008). They further show that, in this context, the lagged error terms can serve as useful instruments in the equations of the transitory shocks. Additional information is thus provided to give consistent estimates of the parameters in the latter equations. Relying on the instrumental variable representation of the BQ model, Pagan and Pesaran show that the weak instrument issue can be solved in cointegrated systems with permanent and transitory disturbances.

Though attractive at first sight, the VEC approach is misleading since nonfundamental representations[3] may easily arise in cointegrating systems (see Blanchard and Quah (1993), and Quah (1995)).

[3] Nonfundamentalness is a major caveat in structural VAR modeling as initially pointed out by Lippi and Reichlin (1993). It refers to situations where the econometrician is less informed than the economist about the true structure of the model. When it occurs, the filtration generated by the vector observable variables X does not correspond to the corresponding natural filtration associated to the vector unobservable structural shocks ε so that the decomposition is no longer unique (see Alessi et al. (2009) for an overview).

To our concern, one major conclusion from what is preceding is that the BQ decomposition will give exactly the same sequence of structural (permanent and transitory) shocks as those obtained by the Cholesky decomposition of the variance-covariance matrix of residuals implied by the Sim's "causal" ordering of variables. When error correction terms are leaved out from the reduced form of the VAR like in (2), the identification of structural innovations can be indifferently done on the basis of one long-run restriction or by use of an equivalent zero constraint on the instantaneous response of one variable of the system. This result still holds in the broader *n*-variate case with *r* cointegrating vectors (see Fisher and Hu (1999)), as well as in the absence of cointegration provided that Wold's ordering is maintained and Granger's long-run causality prevails (see Keating (2009) for a thorough discussion).

Surprisingly, however, this principle of equivalence between the identifying sets of short-run and long-run restrictions has received no attention in the vast empirical literature devoted to the asymmetry in the fixed exchange rate regimes.

3.3 The graphical AS-AD model revisited: long-run or sign restrictions?

In their pioneering work, BE (1992) make use of the textbook graphical version of the AS-AD framework to justify the zero-restriction imposed on the long-run response of output to a transitory shock. If positive, such an impulse can be associated to a displacement of the AD line to right from AD_0 to AD_1 as depicted in figure 1a below. It is followed by an increase in prices (from p_0 to, say, $p_1=(1+\gamma_1)p_0$) like in output (from y_0 to, say, $y_1=(1+\lambda)y_0$). The economy moves along the positively sloped SRAS line in the short run from E_0 to the transitory equilibrium E_1.

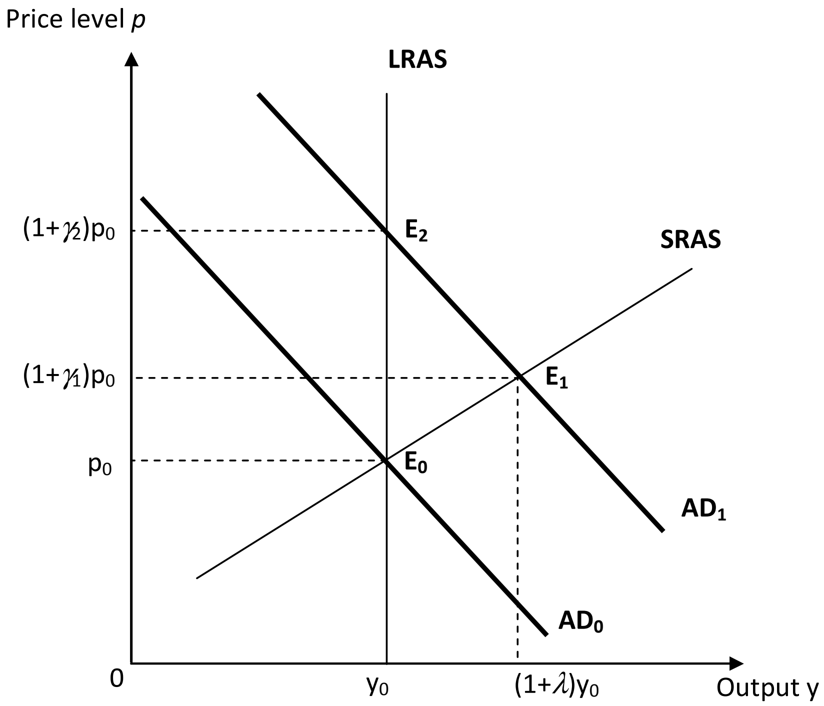

Figure 1a. A positive demand shock in the AS-AD model

As time passes, however, inflation expectations adjust so that AS becomes vertical (see LRAS on the figure) and the economy moves along the new AD_1 line from E_1 to its new long-run equilibrium E_2. The domestic production returns to its natural level y_0 while there is an additional inflationary impact pushing the price level to $p_2=(1+\gamma_2)p_0$.

However this picture may be completed when considering the response of price and output to a permanent shock from the supply side as it can be seen on figure 1b below. If again positive, domestic activity will automatically rise from y_0 to $y_1=(1+\delta_1)y_0$, and finally to reach a higher natural level $y_2=(1+\delta_2)y_0$. Instead, price should fall gradually to $p_1=(1-\delta_1)p_0$. Starting at point F_0 on figure 1b, adjustments in production and prices will continue until the stationary state F_2 is met at the intersection of the AD curve with the new long-run AS line $LRAS_1$.

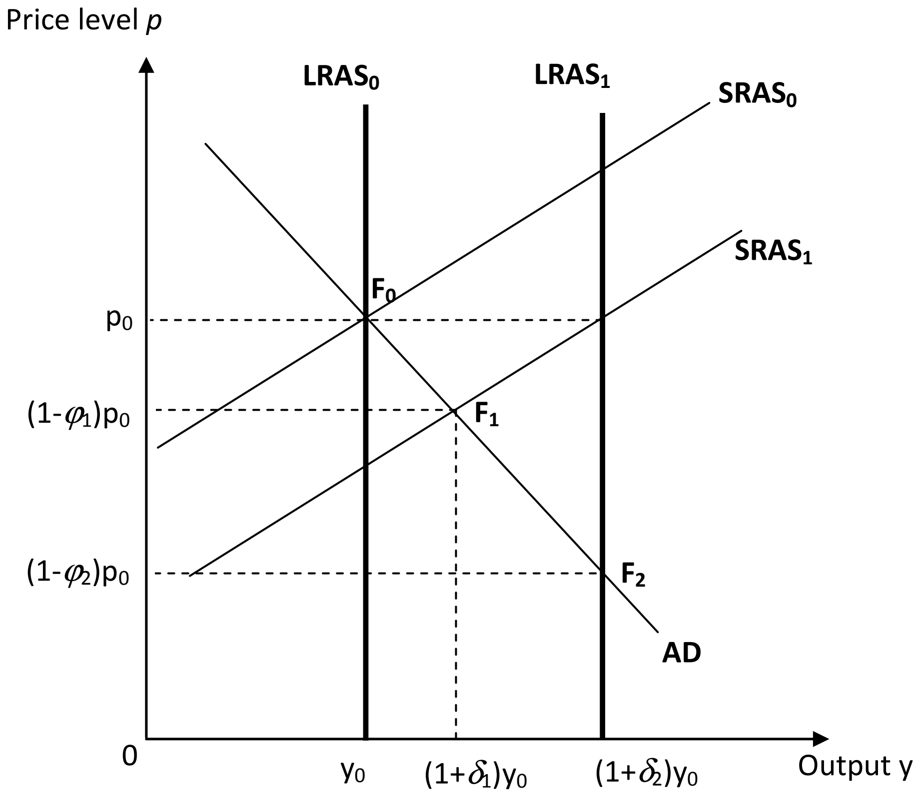

Figure 1b. A positive demand shock in the AS-AD model

All these effects imply specific features of the impulse response functions built from the structural vector autoregression extracted from a reduced form like (1). These are summarized in the following table 1.

Table 1. The cumulative impulse response functions implied by the log-linear AS-AD model

		Cumulative impact on			
		Output (y)		Price level (p)	
		Short run	Long run	Short run	Long run
Type of shock	Transitory (AD curve)	λ	**0**	γ_1	γ_2 ($>\gamma_1$)
	Permanent (AS curve)	δ_1	δ_2 ($>\delta_1$)	$-\varphi_1$	$-\varphi_2$ ($<-\varphi_1$)

Note: All Greek letters refer to positive coefficients.

If the long-run neutrality hypothesis is valid, and provided that the AS and AD relationships are linear, their respective slope coefficients can be recovered as:

$$\eta_{AS} = \frac{\gamma_1}{\lambda}, \text{ and } \eta_{AD} = \frac{\gamma_2 - \gamma_1}{-\lambda}. \tag{26}$$

In practical terms, knowledge of the cumulative IRFs to a transitory shock is sufficient to determine the values of the above ratios. BE (1994) give a nice illustration of the economic meaning of these functions. But they do not go on further to determine precisely the slope parameters. On these grounds, one may be interested in comparing (26) with (19) and (22).

Table 1 also reveals another possible identification strategy for the VAR, inspired by Uhlig's (2005) "agnostic" approach. As it stands, sign-restrictions on the impulse response functions can be easily inferred from this basic macroeconomic setting. Positive demand shocks and cost-push disturbances indeed exert opposite effects on the price level on impact (compare fig. 1a and fig. 1b above). The long-run neutrality constraint may be thus skipped in favor of these non-zero restrictions over a short horizon.

4. Empirical evidence on fixed exchange rate regimes

4.1. Data

Data are taken from the Eurostat database on a monthly basis over the period 1996:01-2008:12. The industrial production index (IPI) in volume is used as a proxy for output. Inflation is measured on the basis of the Harmonized Consumer Price Index (HCPI). All these variables are taken in logs. Price and output are assumed to follow I(1) processes, so that they are first-differenced according to specification (1) above. This approach conforms to what is usually assumed in the empirical literature on shock asymmetry under fixed exchange rate regimes. Few of the past studies indeed run formally unit-root and cointegration-rank tests. Bayoumi and Taylor (1995) is one noticeable exception where Engle and Granger's two-step procedure is applied.

In order to illustrate the previous principles findings and, in particular, to check for the robustness of the empirical findings from the BE approach, we focus on the eleven founders countries which joined the EMU in 1999. Greece is also included to the dataset since its adhesion to the euro was already planned at that time. It also allows for contrast the results for the so-called "PIIGS" with those associated to what was often considered as the core of the former German Mark zone. Taking Germany as the reference country for the EMU allows us to compare our results with BE's initial findings.

4.2. Relative size and paths of permanent and transitory shocks

Table 2 below reports the variance ratios between the identified permanent and transitory shocks in each country over the whole period. Each column refers to a specific procedure of VAR identification. "BQ" refers to Blanchard-Quah's (1989) method, "BE" to Bayoumi-Eichengreen's (1992), "Adjusted-BE" involves the transition matrix as given by formula (14) in the text, and "Choleski" to the well-known approach.

Table 2. Relative size of structural shocks under various decompositions 1996:01-2008:12

	BQ, CEH or Choleski	BE	Adjusted BE	BE (1992)
Germany	**1**	**0.82**	**0.37**	0.82
Austria	**1**	**0.71**	**0.17**	n.a.
Belgium	1	0.44	0.44	1.07
Finland	1	0.22	0.22	n.a.
France	**1**	**0.74**	**0.28**	0.74
Luxemburg	**1**	**0.26**	**0.22**	n.a.
Netherlands	**1**	**0.54**	**0.32**	0.88
Portugal	**1**	**0.96**	**0.22**	0.79
Ireland	**1**	**0.24**	**0.10**	1.62
Italy	1	0.44	0.44	0.91
Greece	1	0.57	0.57	0.53
Spain	1	0.33	0.33	0.68

Note: Figures are ratios of the standard deviations of transitory relative to permanent shocks. The final column reports BE's (1992) initial results in terms of demand relative to supply shocks over 1962-1988.

From table 2, we conclude that non-normalized structural shocks lead systematically to variance ratio less than unity. This means that transitory shocks are smaller than the permanent ones. This result holds whether BE's procedure is corrected by the transition matrix Q or not. The only exception is Portugal where unadjusted BE's approach leads to permanent and temporary shocks of almost equal sizes.

Although all these countries belong to the same currency union, they exhibit markedly differences regarding the shocks which hit their economies. Relying on BE's decomposition, the euro founder members can be divided into two groups: Germany, Austria, and France are characterized by a variance ratio in the [0.7,1[range like Portugal, whereas permanent disturbances dominate by far the transitory shocks in the other Member States.

This picture conforms reasonably well to the core-periphery view of the European Monetary Union. It is broadly consistent with Bayoumi and Eichengreen's (1992) findings for the pre-EMU period (see their reported estimates in the last column of table 2). Contrary to these authors, demand (here transitory) shocks are less sizeable than supply (or permanent) ones as Belgium and Ireland might have experienced in the past decades.

What is also at stake here are the economic consequences of the European monetary unification. One can indeed hardly agree with BE's conjecture that industrial specialization has strengthened in the euro Members States so that demand disturbances now outweigh those from the supply side. Rather, our estimates would give support to the alternative "diversification" hypothesis. The European process seems to be distinct from the one observed at the level of the US regions.

Furthermore, the choice of the transition matrix given by equation (14) may indeed matter for assessing the size of the structural macroeconomic shocks. The corresponding estimates reported in the last column of table 2 reveal to types of countries. The ratio of standard deviations of shocks is left unchanged when modifying BE's procedure in the case of Belgium, Finland, Italy, Greece, and Spain. This contrasts with the sharp decrease experienced by the remaining countries under study. The switch from the original BE's method to the proposed decomposition given by the identity (12) implies a further reduction in the size of the transitory shocks relative to the permanent component.

Adjusting BE's factorization for the transition matrix Q may modify one's view about the way EMU actually operates. From the third column of table 2, it is uneasy to distinguish the core from the periphery of the euro area on the sole basis of the size of domestic shocks.

4.3. Shock asymmetries and the European currency union

Let us now consider the sensitivity of asymmetry measures to the set of identifying restrictions. We first consider the correlation coefficients between each type, permanent or transitory, shock. All of them are computed against Germany. Correlations between permanent shocks are reported in table 3, those associated to temporary innovations can be found in table 4 below.

Table 3. Correlation coefficients of permanent shocks (against Germany, 1996:01-2008:12)

Country	**BQ**	**BE**	**Adjusted BE**	**CEH**	**Choleski**	BE (1992)
Austria	0.215	0.215	0.231	0.200	0.231	n.a.
Belgium	0.315	0.315	0.317	0.234	0.316	0.61
Finland	0.335	0.335	0.352	0.256	0.356	n.a.
France	0.333	0.333	0.338	0.353	0.340	0.54
Luxemburg	0.133	0.133	0.115	0.155	0.108	n.a.
Netherlands	0.137	0.137	0.134	0.122	0.131	0.59
Portugal	0.249	0.249	0.274	0.280	0.279	0.21
Ireland	0.230	0.230	0.234	0.193	0.238	0.06
Italy	0.432	0.432	0.437	0.364	0.436	0.23
Greece	0.172	0.172	0.189	0.193	0.194	0.14
Spain	0.376	0.376	0.384	0.232	0.384	0.31

Note: The final column reports BE's (1992) initial results in terms of supply shocks over 1962-1988.

Results from table 3 illustrate the equivalence principle between BQ's and the Choleski decomposition scheme as demonstrated by Ribba (1997), and Fisher and Hu (1999, 2000). A similar conclusion can be drawn from table 4 below. Because they give similar series of structural shocks country by country, the choice between these two particular sets of long- and short-run restrictions is inconsequential for the correlation coefficients themselves. Therefore, referring to the AS-AD framework in order to assume the long-run neutrality of output to transitory shocks does not matter for the appraisal of stochastic asymmetry within a currency union like the euro area. This new empirical evidence, jointly with the theoretical results, directly challenge the common econometric practice inherited from the influential works of Bayoumi and Eichengreen in that field.

Estimated values reported on table 3 confirm that the measurement of the correlation between permanent shocks does not depend to the way of factorizing the covariance matrix of the VAR residuals. Assigning unit variance to all shocks – like in Blanchard and Quah (1989) – or allowing for structural disturbances of unequal sizes – as suggested by Bayoumi and Eichengreen (1992, 1994) – leads to the same level of asymmetry in terms of permanent shocks. This is well in accordance with BE's premise: their departure to the BQ approach should imply just a rescaling of shocks, thereby leaving their other properties unchanged.

While BE put the emphasis on the discrepancies between the core and the peripheral countries during the pre-EMU phase, greater homogeneity is found amid the euro founder Members during 1996-2008. As shown on table 3, asymmetry in terms of permanent shocks has increased in the core (Belgium, France, and the Netherlands) against Germany. At the opposite, permanent shocks to the periphery (namely the PIIGS) seem to be more correlated to those hitting the German economy. Based on this criterion, Greece is as far to the euro area as other small open countries like Ireland, the Luxemburg or even the Netherlands.

Table 4 below gives the corresponding correlation estimates between transitory shocks over the whole sample period. The equivalence principle between the BQ the Choleski factorizations is still valid. However, there is much variability among the correlation values. These are negative, though close to zero, for Greece and the Netherlands. They are of the same order of magnitude as the asymmetry in permanent shocks only in Austria and France. Shocks to the remaining countries are essentially idiosyncratic when they have temporary effects on domestic output. This is more in accordance with the core-periphery view, though it deserves some words of caution as in Bayoumi and Eichengreen (1992) (see their own estimates in the last column of table 4).

Table 4. Correlation coefficients of transitory shocks (against Germany, 1996:01-2008:12)

Country	**BQ**	**BE**	**Adjusted BE**	**CEH**	**Choleski**	BE (1992)
Austria	0.335	*-0.184*	0.175	0.232	0.352	n.a.
Belgium	0.125	*-0.107*	0.022	0.291	0.158	0.33
Finland	0.091	0.229	0.214	0.241	0.093	n.a.
France	0.340	0.349	0.350	0.266	0.335	0.35
Luxemburg	0.116	*-0.013*	0.023	0.137	0.148	n.a.
Netherlands	*-0.096*	0.084	*-0.030*	0.166	*-0.102*	0.17
Portugal	0.061	0.289	0.213	0.224	0.030	0.21
Ireland	0.177	0.207	*-0.031*	0.169	0.203	0.08
Italy	0.062	*-0.020*	0.023	0.348	0.068	0.17
Greece	*-0.086*	0.015	*-0.044*	0.005	*-0.099*	0.19
Spain	0.103	*-0.104*	0.037	0.240	0.124	0.07

Note: The final column reports BE's (1992) initial results in terms of demand shocks over 1962-1988.

But things turn to be very different if one follows BE’s methodology. The second column of table 4 indeed reveals that the picture about asymmetry in terms of transitory shocks is modified. In almost all cases, correlations change of magnitude if we switch from the BQ factorization to the (unadjusted) BE decomposition. For example, the correlation between the Irish and the German transitory shock rises by a third roughly. It doubles at least in Finland, and even quadruples in the Portuguese case.

If the BQ approach were viewed as the relevant one, asymmetry in the temporary “surprises” would then be underestimated. However, we are led to the opposite conclusion as concerns Austria, Belgium, Italy, and Spain: there is now evidence of strong asymmetries since correlations between BE shocks turn out to be negative. There are only two Member States – namely, France and Ireland – whose results are unaffected.

Although things remain the same in terms permanent shocks, the situation is now completely different, if not reversed, when considering transitory disturbances. It is thus no longer possible to conclude with Bayoumi and Eichengreen that relaxing the assumption of equal and unitary variances would just amount to a rescaling of the innovations of the SVAR.

As demonstrated in section 2, the BE factorization is actually defined up to some orthonormal matrix. In particular, a transition matrix Q can be found so that the structural shocks we get by the "adjusted" BE decomposition behave like those corresponding to the BQ procedure. This should imply similar correlation coefficients. Even though it is the case for permanent shocks (table 3), there are noticeable discrepancies as concerns the transitory components (table 4). The estimated value is reduced by one half or more in Austria, Belgium, Luxemburg, Italy, and Spain. Taken in absolute values, it doubles or more in Finland, the Netherlands, Portugal, and Greece. The equivalence prevails in France only. By contrast, a change in the sign of the correlation coefficient is observed in the Irish case if the BQ result is taken as a benchmark.

Though surprising at first sight, these results can be explained by the non-uniqueness of the transition matrix as it has already been stressed in section 2. In principle, one has to pick up one of the eight possible writings of Q given by equation (12). It is therefore easy to recover a positive correlation for Ireland by an appropriate transformation of Q. Still, none of the available transition matrices enables to retrieve exactly the BQ-type correlation coefficients between the transitory shocks in most cases.

The difficulty to retrieve the BQ correlations from the "adjusted" BE factorization may lie in the estimated transition matrices. Table 5 reports the Q matrices used to built the tables 2 to 4. It is worth highlighting that Q is close to the identity matrix in the vast majority of cases. Off-diagonal elements seem to be highly sensitive to, even small, departures from unity on the principal diagonal of this type of rotation matrix. But the cases of Austria and Portugal are left unexplained.

4.4. Responses to shocks and the underlying AS-AD model

The evidence about slope estimates of the aggregate supply function is rather mixed when the CEH identification strategy is employed. Table 6 below reports the corresponding figures for the founder members of the euro area during 1996-2008. According to the above system (15), the slope parameter for the AS curve is given by $1/\alpha$. Positive excepted values are reported in bold face. For comparison purposes, the results obtained by Lee and Crowley (2010) for the same group of euro Members are shown in the last column of this table. These come from a New Keynesian model augmented by a Taylor rule followed by the ECB.

Table 6. Slope estimates of AS curves under CEH identifying restrictions (1996:01-2008:12)

Lag-order p of the VAR	1	2	3	4	5	6	7	8	9	10	11	12	**LC (2010) estimates**
Germany	-0.27	-0.16	-0.19	-0.58	-0.78	-2.38	**0.70**	**0.58**	**0.76**	**0.63**	**0.87**	-0.04	**0.02**
Austria	**0.01**	-0.04	-1.41	-0.16	**0.07**	**0.03**	**0.02**	**0.04**	**0.08**	**0.04**	-0.02	-0.10	**0.10**
Belgium	**5.56**	**6.67**	-3.85	-3.70	-25,00	**0.03**	**0.16**	**0.25**	**0.26**	**0.27**	**0.32**	**0,00**	**0.26**
Finland	**33.33**	**0.15**	**0.24**	**0.50**	**0.38**	**0.18**	**0.04**	**0.03**	**0.13**	**0.16**	**0.06**	-0.12	**0.05**
France	**0.28**	**0.05**	**0.05**	-0.23	-1.10	-0.30	**0.04**	**0.04**	**0.08**	**0.04**	-0.03	-0.11	**0.23**
Luxemburg	-1.49	-1.09	-0.81	-1.11	-3.13	**50.00**	**0.13**	**0.16**	**0.30**	**0.12**	**0.10**	**0.01**	**0.05**
Netherlands	**1.37**	**0.49**	-1.12	-0.71	-2.70	**0.50**	**0.03**	**0.71**	**0.56**	**0.51**	-0.09	-0.44	**0.13**
Portugal	**0.44**	**0.10**	-0.08	-0.20	-0.07	**0.02**	-0.08	-0.21	-0.47	-0.26	-0.27	**0.10**	**0.06**
Ireland	-0.37	-0.22	-0.09	-0.10	-0.25	**1.19**	**0.04**	-0.20	-0.23	-0.49	**0.00**	**1.85**	**0.07**
Italy	-684.93	-1.32	-1.12	-2.27	-2.63	**0.30**	**0.24**	**0.36**	**0.36**	**0.54**	**0.46**	-0.04	**0.13**
Greece	**0.37**	-11.11	-5.00	-4.55	-5.00	-0.63	-0.83	**5.26**	**1.59**	**1.43**	**2.08**	**0.18**	**0.47**
Spain	-3.85	-0.27	**50.00**	-0.59	-1.03	-0.68	-0.12	-0.09	**0.23**	**0.24**	**0.03**	-0.21	**0.14**

The slope estimates with the CEH approach exhibit considerable variability with the chosen lag-order of the VAR system. Adding just one more lag to the dependent variables may lead to either a sudden change in the order of magnitude or to a sign reversal, as it is observed in all the countries under study. As concerns Austria, the parameter α varies from -26.32 to 151.29 when it is computed as in (19). Nine lags in the vector autoregression give the least unreasonable value of 12.2. This leads to an AS slope coefficient of 0.08 close to Lee and

Crowley's (LC, 2010) result. There is thus evidence of relatively flat AS curves in the founders of the euro area.

Discrepancies are observed between the CEH SVAR approach and the LC New Keynesian model. These are particularly sharp in the German case, our reference country for the bilateral comparisons. More seriously, it appears from table 6 that the German Phillips curve is steeper than in France, contrary to the empirical evidence from New Keynesian DGSE models (e.g. Brissimis and Skotida (2008) among others).

This may be explained by the ECB's commitment to an interest rate policy rule which is accounted for in these general equilibrium models. In addition, the slope coefficient of the AD curve is left unconstrained. It proves to be systematically lower than unity and to vary amid the euro Member States. Lee and Crowley (*o. p.*) reports values ranging from 0.01 to 0.21. This is inconsistent with the full price indexation hypothesis made by CEH (2006). If the bivariate VAR setting is misspecified, there may well be strong bias in the parameter estimates as well as in the impulse response functions (see Braun and Mittnick, 1993). This is a crucial issue since α comes from the long-run dynamic multipliers.

Table 5 also shows that implausible (negative) values of the AS slope are usually obtained with small lags in the vector autoregression. Estimates of α are also much more sensitive to the choice of a low value of p. This may question the relevant choice of the lag-order of the VAR process. Worrying about parsimony, the econometrician often relies on standard information criteria, especially Schwartz's conservative one, to get an "optimal" value for p. The "best" value is often 1, rarely 2, as it is the case in the estimated VARs underlying the building of tables 2 to 4. But, as it is apparent here, this choice may be viewed as too conservative if one follows the CEH procedure of VAR identification. Unreliable estimates of the slope of the AS curve would then be obtained.

As already pointed out by Braun and Mittnick (1993), adding lags to the autoregressive component of the dynamic system may circumvent (at least part of) the misspecification problems to recover the true impulse response functions. As regards the CEH approach, it may be reflected in a severe biased estimate of the slope parameter of the AS curve. This may be due to omitted moving average terms which often appear in a New Keynesian framework under the rational expectation hypothesis. They are clearly neglected in “pure” VAR reduced forms like (1).

Slope estimates based on the graphical representation of the AS-AD model are reported on with each other: the BQ strategy as depicted in table 1 and Uhlig’s pure sign approach. According to the first method, the estimated coefficients have the expected sign in almost all cases. A major exception is Netherlands for which the identified shocks can hardly be interpreted as supply and demand disturbances because of the complete sign reversal in their observed effects on output and prices. As concerns Belgium, its AD curve seems also to be positively sloped, contrary to what the inflation–unemployment tradeoff would have implied. Another striking feature is that the strong heterogeneity in the implied slope values when the long-run neutrality of output is assumed. AD curves are generally found to be steeper than (short-run) AS curves, reaching unrealistic levels in the core (Austria and Germany) like at the periphery of the euro area (Greece and Ireland). AS slopes are (more or less) in line with the findings of other recent studies (see last column of table 6 above). Instead, AD slope parameters exhibit strong discrepancies with Lee and Crowley’s (2010) values (see the last column of table 7).

Table 7. Slope estimates from the graphical view of the AS-AD model (1996:01-2008:12)

	Choleski decomposition		**Sign restrictions**				**LC (2010) estimates**
			From **minimum** response		From **maximum** response		
	AS line	**AD** line	**AS** line	**AD** line	**AS** line	**AD** line	**AD** line
Germany	0.42	-5.00	0.40	-0.06	0.16	-0.03	**-0.11**
Austria	0.74	-14.29	0.00	-6.51	0.17	-0.03	**-0.01**
Belgium	0.05	0.18	1.01	n.d	0.42	0.13	**-0.07**
Finland	0.04	-0.31	0.51	n.d.	0.19	0.03	**-0.10**
France	0.10	-0.56	0.23	-0.21	0.10	n.d.	**-0.04**
Luxemburg	0.06	-1.64	0.40	-3.25	0.21	-0.27	**-0.20**
Netherlands	-0.41	0.73	0.05	-2.29	0.22	-0.05	**-0.02**
Portugal	0.10	-1.04	0.08	-0.40	0.00	-0.49	**-0.04**
Ireland	0.24	-12.50	0.07	-0.32	0.09	0.01	**-0.14**
Italy	0.00	0.00	0.13	1.60	0.02	n.d.	**-0.12**
Greece	0.01	-25.00	0.04	-0.51	0.16	n.d.	**-0.21**
Spain	0.09	-0.26	0.33	-0.05	0.05	n.d.	**-0.03**

Note: Undetermined values of slope parameters are abbreviated with n.d..

If we switch to the agnostic approach, results differ markedly. Remember that the estimates of slope coefficients are now obtained from impulse response functions based on the VAR in levels. The values given by the pure sign approach are computed according to formulas in table 1 from the impulse response functions shown on graphics 2 and 3 in the annex. Sign restrictions were imposed output and price responses during the next 3 months following a transitory (demand) shock. 500,000 simulations have been launched of which at most 50,000 “successes” have been collected. The range of effects is revealed by the minimum and maximum impacts on each of these aggregates over a five-year horizon (or equivalently 60 months). For ease of comparisons, the response functions derived from the Choleski decomposition are also reported on these graphics.

The estimates for AS curves are close to those reported on table 6 in a majority of countries. As emphasized by the previous studies, AS like AD curves are very flat. Our results do not give support to the full price indexation assumption made by CEH (2006), even though a statistical test cannot be put formally. Table 7 also shows the difficulties in calculating the slope the aggregate demand relationship. These are sometimes impossible to determine or wrongly signed because the impulse response function of industrial production to a transitory shock does not conform to what is expected from the AS-AD model.

5. Conclusion

This paper has discussed the relevance of long-run restrictions in structural VAR models within the textbook AS-AD theoretical framework. As popularized by Bayoumi and Eichengreen (1992, 1994), the latter is often used as the economic background to investigate the empirical properties of shocks under alternative exchange rate agreements. Our contribution in this field is twofold.

As regards structural VAR modeling, it is shown how Blanchard and Quah's (1989) stratregy is linked to its competing alternatives in order to distinguish permanent from transitory shocks. In particular, it is shown how the modified procedure suggested by Bayoumi and Eichengreen themselves may depart significantly from BQ's. However, a transition matrix can be found to back out BQ's decomposition of the VAR residuals. Still, this particular rotation matrix is not unique which adds to the identification problem. Furthermore, relaxing auxiliary assumptions in VAR identification – especially the orthogonalization of shocks in a given country – may lead to significant departure from the BQ decomposition scheme. Since VAR identification through long-run restrictions has been severely questioned, short-run alternatives have also been considered here. We are thus led to emphasize a important result which has been disregarded by the empirical literature of fixed exchange rate regimes: zero-restrictions on either long-run or comtemporaneous responses of variables to shocks may be strictly equivalent. As such, a Choleski decomposition is not a alternative to BQ's approach.

These new insights in structural VAR modeling have important consequences for the empirical analysis of shock asymmetry under a fixed exchange rate regime. Our previous findings have been illustrated the experience of the eleven founder members of the euro area (plus Greece) during 1996-2008.

Taking into account the transition matrix from BE to BQ decompositions matters for evaluating the relative size of permanent relative to transitory shocks. Though permanent shocks always dominate, the country ranking appears to be very sensitive to inclusion of the transition matrix to identify both sources of structural shocks. The updated evidence provided also clearly conflicts with BE's premise that the currency union would have fostered industrial specialization thereby increasing the relative size of transitory (demand) disturbances. Furthermore, linking the BE decomposition to the BQ one through the transition seems to be inconsequential for the measurement of asymmetry in permanent shocks, whereas it has a dramatic influence on the empirical assessment of asymmetry in the transitory component. It is also shown that the issue raised by BE's identification strategy is further complicated by the non-uniqueness of the transition matrix itself. The former is indeed only defined up to a given rotation. This issue is similar to the one pertaining to the Given's matrices underlying the (short-run) sign restrictions for the VAR identification. From this perspective, the basic AS-AD diagram used by BE may help recover the slope coefficients of the AS and AD curves. However, sign-restrictions according to Uhlig's (2005) pure agnostic approach give in general more reliable estimates of these slope parameters than zero-constraints on the response functions derived from VAR estimates do. Aggregate demand

and well supply curves are usually found to be flat, but they differ substantially from one euro Member State to another.

From this perspective, the conclusions drawn from our analysis may also have implications to other important economic issues. Beyond the AS-AD diagram, similar concerns about structural VAR modeling can indeed be found in the business cycle literature where some knowledge of the underlying dynamic macroeconomic setting is needed (Canova (2009) for a discussion about dynamic stochastic general equilibrium (DSGE) models). It may also matters for the identification of technological shocks as well as in the analysis of the monetary policy, taking the long-run (cointegrating) relationships among the economic variables into account (Pagan and Pesaran (2008)). Though promising, VAR identification is still an issue.

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Annex

Table 5. Factorizations of the covariance matrix of the VAR residuals and the transition matrix between the BE and BQ decompositions

	Transition matrix Q			BQ factorization C_{BQ}			BE factorization C_{BE}			Choleski decomposition	
Germany	0.862	-0.506		0.012	-0.003		0.801	-0.598		0.012	0
	0.506	0.862		0.001	0.003		0.627	0.779		1.0e-04	0.003
Austria	0.862	-0.506		0.015	0.002		0.793	0.609		0.015	0
	0.506	0.862		-1.7e-04	0.003		-0.565	0.825		1.3e-04	0.002
Belgium	0.999	-0.029		0.017	3.7e-04		0.999	0.05		0.017	0
	0.029	0.999		-1.6e-04	0.007		-0.051	0.999		-1.4e-06	0.007
Finland	0.999	0.003		0.019	-1.5e-05		0.999	-0.004		0.019	0
	-0.003	0.999		5.4e-04	0.004		0.155	0.988		5.4e-04	0.004
France	0.898	0.44		0.011	-0.002		0.814	-0.581		0.011	0
	-0.44	0.898		4.4e-04	0.003		0.584	0.812		8.2e-06	0.003
Luxemburg	0.997	0.082		0.029	-0.001		0.995	-0.104		0.029	0
	-0.082	0.997		6.2e-04	0.006		0.183	0.983		4.9e-04	0.006
Netherlands	0.964	0.264		0.018	-0.002		0.93	-0.367		0.018	0
	-0.264	0.964		9.4e-04	0.005		0.446	0.895		4.2e-04	0.005
Portugal	0.825	0.566		0.024	-0.003		0.736	-0.677		0.024	0
	-0.566	0.825		6.2e-04	0.004		0.696	0.718		9.6e-05	0.004
Ireland	0.958	0.286		0.049	-0.001		0.95	-0.313		0.049	0
	-0.286	0.958		-2.9e-04	0.004		0.221	0.975		-4.2e-04	0.004
Italy	0.999	-2.2e-04		0.009	1.6e-06		1	0		0.009	0
	2.2e-04	0.999		-1.8e-04	0.004		-0.044	0.999		-1.8e-04	0.004
Greece	0.999	0.017		0.021	-4.7e-04		0.999	-0.039		0.021	0
	-0.017	0.999		8.1e-05	0.012		0.024	0.999		-1.8e-04	0.012
Spain	0.998	-0.06		0.013	3.9e-04		0.996	0.09		0.013	0
	0.06	0.998		-6.4e-06	0.004		-0.062	0.998		1.2e-04	0.004

Figure 2. Response of output to a transitory shock: Choleski decomposition *versus* pure sign approach

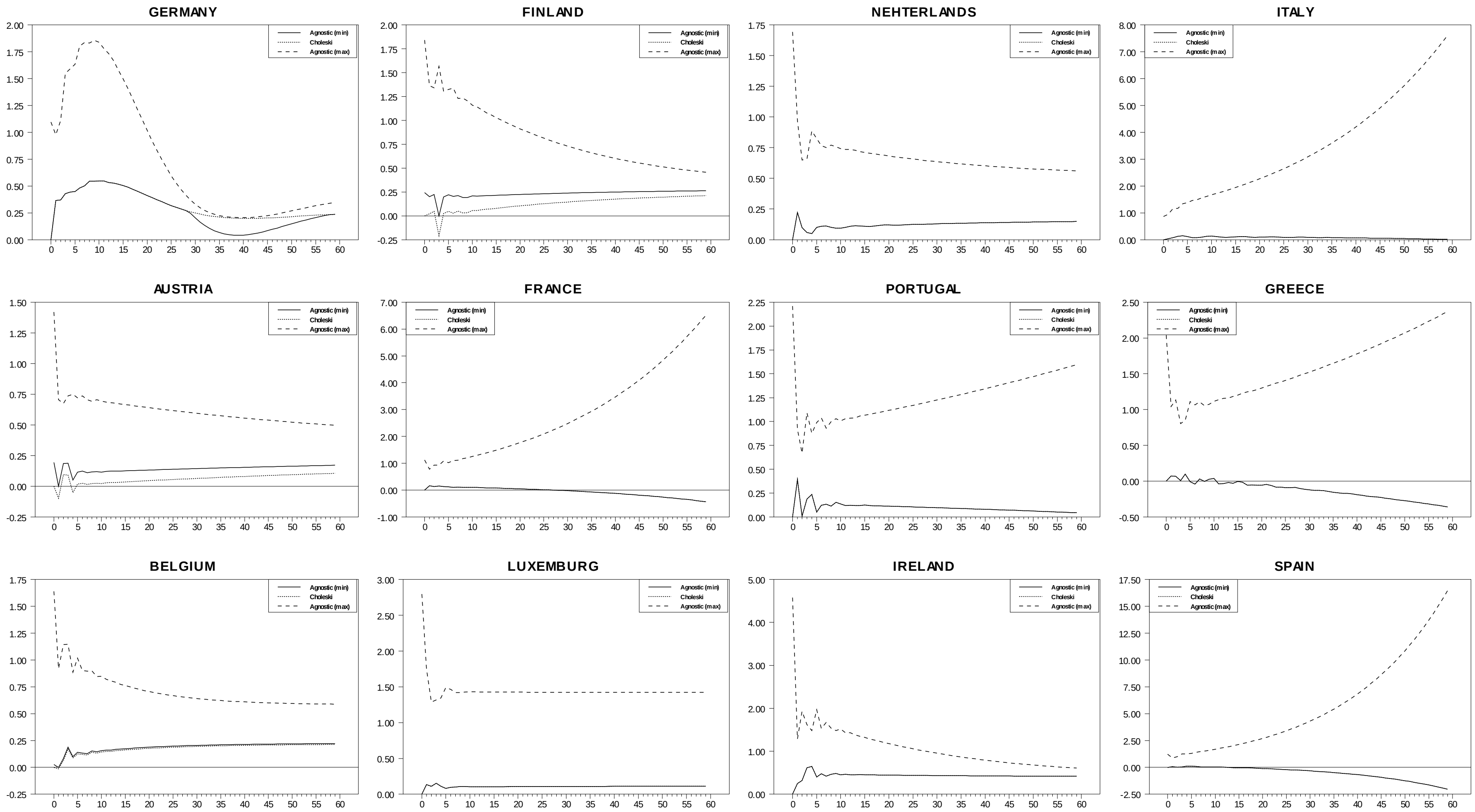

Figure 3. Response of HCPI to a transitory shock: Choleski decomposition *versus* pure sign approach