
NOTES D'ÉTUDES

ET DE RECHERCHE

CONVERGENCE IN HOUSEHOLD

CREDIT DEMAND

ACROSS EURO AREA COUNTRIES:

EVIDENCE FROM PANEL DATA

Olivier de Bandt, Catherine Bruneau, Widad El Amri

October 2006

NER - R # 158



**CONVERGENCE IN HOUSEHOLD
CREDIT DEMAND
ACROSS EURO AREA COUNTRIES:
EVIDENCE FROM PANEL DATA**

Olivier de Bandt, Catherine Bruneau, Widad El Amri

October 2006

NER - R # 158

Les Notes d'Études et de Recherche reflètent les idées personnelles de leurs auteurs et n'expriment pas nécessairement la position de la Banque de France. Ce document est disponible sur le site internet de la Banque de France « www.banque-france.fr ».

Working Papers reflect the opinions of the authors and do not necessarily express the views of the Banque de France. This document is available on the Banque de France Website "www.banque-france.fr".

Convergence in Household Credit Demand Across Euro Area Countries: Evidence from Panel Data*

O. de Bandt[†], C. Bruneau[‡], W. El Amri[§]

(This Version: October 2006)

*Any views expressed in the paper are the authors' own and do not necessarily reflect those of the Banque de France or the Eurosystem. They thank Claudine Guibert for providing excellent research assistance in compiling the database, as well as seminar participants at Banque de France, Bank of England and the Maastricht University for useful comments.

[†]Banque de France, corresponding author. E-mail olivier.debandt@banque-france.fr, address: Banque de France, 46-1405 DAMEP, 39 rue croix des petits champs, 75049 Paris Cedex 01, phone +33 1 42 92 28 80, Fax + 33 1 42 92 49 50.

[‡]Banque de France and University of Paris X, EconomiX.

[§]Banque de France and University of Paris X, EconomiX.

Abstract

The paper contributes to the literature on the convergence of financial systems in the euro area by estimating household credit demand in individual countries. Using the ARDL framework advocated notably by Pesaran *et al.* (1999), the paper provides evidence on the convergence of long run credit demand determinants (interest rates, investment and house prices) among the largest euro area countries, while short run dynamics remain heterogeneous across countries. The paper also demonstrates that the equation uncovers demand rather than supply behaviour.

Keywords : Credit demand, panel cointegration, households, bank profitability.

JEL classification: E51, C31, C32, C33.

Résumé

L'article contribue à la littérature sur la convergence des systèmes financiers dans la zone euro en estimant une équation de demande de crédit des ménages dans différents pays. En utilisant le modèle ARDL proposé en particulier par Pesaran *et al.* (1999), l'article met en évidence la convergence des déterminants à long terme de la demande de crédit (taux d'intérêt, investissement, prix immobiliers) au sein des plus grands pays, alors que les dynamiques de court terme demeurent hétérogènes. L'article vérifie aussi que l'équation correspond à un comportement de demande plutôt que d'offre de crédit.

Mots-clés : Demande de crédit, cointégration en panel, ménages, profitabilité bancaire.

Classification JEL : E51, C31, C32, C33.

Non technical summary

The paper studies the convergence of financial systems in the euro area by focusing on household credit demand. Two questions are particularly addressed, first of all, we investigate whether credit markets remain heterogenous across euro area countries or have become more similar, as a consequence of the Single Capital Market programme of the early 1990s and EMU from 1999 onwards; second, we look for reliable estimates of key parameters of credit demand like the semi-elasticity of credit to interest rates, as well as the impact of house prices on credit demand, against the background of sustained increase since the mid 1990s but also divergent situations across countries (Germany experiencing, on the contrary, subdued increase for both house prices and credit).

From the methodological point of view, while studies usually rely on individual time series, the paper considers the dynamics of a panel of euro area countries, using panel data techniques with the view to increasing the sample size in particular because credit data are often subject to methodological breaks, hence reducing the availability of long time series. The paper implements a version of the AutoRegressive Distributed Lag (ARDL) model advocated by Pesaran *et al.* (1999), which jointly allows estimating a long run equation and its impact on the short run dynamics of a variable of interest. Moreover, ARDL models firstly introduced to deal with time series has been easily extended to panel data. In the latter case, it is possible to impose long run homogeneity with possibly heterogeneous short run dynamics across individuals. This appears to be a convenient characterisation of credit dynamics in the euro area, where long run behaviour is expected to be common across countries as it is mainly determined by economic behaviour. Short run dynamics, however, may remain divergent as they express the permanence of differences across countries, pertaining to idiosyncrasies in institutions, languages or culture. In this paper, we examine whether the Pooled Mean Group (PMG) specification, which imposes long run homogeneity and leaves the short run dynamics unconstrained, is supported by the data. Such a model is compared to standard ECM models in section 3.

One drawback of the analysis, however, is that the the PMG model assumes the existence of cointegration. In order to test such an hypothesis, the paper implements two different types of tests, either Kao's (1999) test, which is based on the pooled panel, or on in individual time series, as for the Pesaran *et al*'s (2001) "bounds tests". These tests are both discussed in section 4.

The empirical results are based on quarterly data for the period 1991:4-2005:4 et cover 9 countries (Luxembourg, Portugal and Greece are always excluded, due to the lack of sufficiently long time series). The paper validates the existence of a long run equation between credit volume, housing investment, long run nominal interest rate and a relative house-price index. In addition, we interpret this equation as a demand equation, as the corresponding estimated residuals are found not to be correlated with supply factors like bank profitability. This equation can be constrained to be homogeneous in the long run for 7 European countries (*i.e.* after exclusion of Belgium and Austria) and cointegration is validated for the 5 largest countries (Germany, Spain, France, Italy and the Netherlands). Specification tests and dynamic simulations indicate a reasonable fit, while the results appear to be stable over time.

The final result is that the PMG specification is supported by the data, so that we can conclude that there is evidence of common long term economic determinants of household credit demand, at least in the largest euro area countries, although short run adjustments remain different.

Résumé non technique

L'article étudie la convergence des structures financières dans la zone euro en se concentrant sur la demande de crédit des ménages. Deux questions sont notamment abordées, tout d'abord, nous étudions si les marchés du crédit demeurent hétérogènes entre les pays ou si leur modes de fonctionnement sont devenus plus similaires à la suite de la mise en place du Marché Unique dans l'Union Européenne au début des années 1990 ou de l'Union Monétaire à partir de 1999. Ensuite, nous cherchons à obtenir des estimateurs fiables des paramètres clés de la demande de crédit, comme par exemple la semi-élasticité des crédits aux taux d'intérêt ou l'impact des prix immobiliers, dans le contexte de hausse continue des prix du logement depuis le milieu des années 1990, malgré des situations différentes entre les pays (l'Allemagne faisant au contraire l'expérience d'une faible hausse à la fois des prix immobiliers et du crédit aux ménages).

Sur le plan méthodologique, alors que la plupart des études portent sur des données chronologiques sur des pays pris individuellement, l'article considère la dynamique d'un panel de pays de la zone euro, en mobilisant des techniques d'analyse des données de panel dans le but d'accroître la taille de l'échantillon, dans le mesure où les statistiques sur les crédits font souvent l'objet de ruptures méthodologiques, réduisant par là la longueur des séries disponibles. Le papier met en oeuvre une version du modèle autorégressif à retards échelonnés (en anglais: ARDL) proposé par Pesaran *et al.* (1999), qui permet d'estimer une relation de long terme et de caractériser conjointement l'impact de cette relation dans la dynamique de court terme d'une variable d'intérêt. L'estimation d'un tel modèle étudiée d'abord dans le cadre des séries temporelles a été étendue sans difficulté à des données de panel. Dans ce cas, il est de plus possible de tester l'homogénéité de la relation de long terme et/ou de la dynamique de court-terme pour les différents individus. Ce type de test est intéressant pour étudier la dynamique du crédit dans la zone euro. On peut en particulier examiner si les déterminants de long terme sont identiques entre les pays, ce qui est *a priori* attendu, parce qu'ils correspondent à des comportements économiques homogènes dans les différents pays de la zone euro. Conjointement, on peut vérifier que la dynamique de court terme reste divergente entre les pays, en raison de la permanence de spécificités nationales, que l'on peut relier à des effets institutionnels, linguistiques ou culturels. La démarche empirique consiste donc à vérifier si la spécification Pool Mean Group (PMG) - homogénéité à long-terme et hétérogénéité à court terme - est en accord avec les données disponibles. Un tel modèle est comparé au modèle à correction d'erreur standard dans la section 3.

Une des limites de l'analyse est cependant que le modèle PMG fait l'hypothèse que les variables sont cointégrées pour chaque individu. Afin de tester cette hypothèse, le papier met en oeuvre deux types de tests, à savoir le test de Kao (1999), qui est fondé sur des données empilées mais avec effets fixes, et les "bounds tests" proposés par Pesaran *et al.* (2001) sur données chronologiques individuelles. Ces deux types de tests sont discutés en section 4.

L'analyse empirique est menée sur des données trimestrielles portant sur la période 1991:4-2005:4 et sur 9 pays (le Luxembourg, le Portugal et la Grèce sont exclus, du fait de l'absence de séries suffisamment longues). Le papier valide l'existence d'une relation de long terme entre le volume de crédit, l'investissement-logement, le taux d'intérêt nominal à long terme et un indicateur de prix relatif du logement. Nous interprétons cette équation comme une équation de demande de crédit, dans la mesure où les résidus estimés ne sont pas corrélés avec des indicateurs d'offre de crédit, comme la profitabilité des banques. L'homogénéité de l'équation de long terme peut

être validée pour 7 pays européens (c'est-à-dire après exclusion de l'Autriche et de la Belgique) et être assimilée à une relation de cointégration pour les 5 plus grands pays (Allemagne, Espagne, France, Italie et Pays-Bas). Les tests de spécification et les simulations dynamiques indiquent que le modèle présente une bonne qualité d'adéquation aux données. De même le modèle apparaît stable dans le temps.

Au total, la spécification PMG est conforme aux données, et nous pouvons conclure que les structures financières des pays de la zone euro évoluent dans le sens de la convergence des comportements économiques, puisque la dynamique de long terme de demande de crédit est commune, malgré la divergence des ajustements de court terme.

1 Introduction

In the euro area, credit dynamics play an important role in the transmission of the Single Monetary Policy. As opposed to the “money view”, where the transmission channels of monetary policy is based on the direct effect of interest changes on household and companies spending, the “credit view” argues that if financial markets are incomplete or imperfect, it is important to consider the availability of external credit, which may amplify the previous transmission channels of monetary policy (Bernanke, 1988). This is all the more important in the euro area since banks are the main providers of funds to households and companies (ECB, 2002). In addition, while money markets are harmonized, credit markets remain, to a certain extent, segmented, due to differences in language, institutions and competitive environment (see Neven and Röllner, 1999). The paper investigates household credit demand in the euro area and the extent to which it displays similar patterns across countries. Indeed, it is interesting to assess whether the introduction of EMU has changed the functioning of credit markets. Special attention is also devoted to the response of credit demand to house prices which have significantly increased in many countries since the mid 1990s.

An obvious limitation to the traditional time series approach applied to the analysis of credit markets (Friedman and Kuttner, 1993, Fase, 1995) is the lack of long time series of observations. Extending the analysis to a panel of euro area countries may overcome such difficulties and we rely on recent papers in international economics that have addressed the issue of cross-country comparisons, with a view to accommodating the heterogeneity between individuals/countries in panels.

We estimate a long run relationship which characterizes the relationship between household credit in the euro area (in real terms) and investment, interest rate, as well as house prices as fundamentals. The analysis focuses on 9 euro area countries for which sufficiently long quarterly time series are available. We examine whether it is possible to find an homogeneous long run relationship, with common coefficients for all countries, but we allow for specific short run dynamics, using the ARDL (AutoRegressive Distributed Lags) model proposed by Pesaran *et al.* (1999). The intuition is that short run dynamics is more likely to be affected by institutional factors that differ across countries, while long run evolutions are driven by similar economic determinants. Anticipating on the results, the paper shows that household credit variable, in euro area countries, exhibit similar patterns in the long run, but that it can only be interpreted as a cointegration relationship in the largest countries, while the stability of short run dynamics indicates that differences persist over the sample period.

Such a class of models describes the dynamics of a series according to a single equation approach, where the change in the selected endogenous variable (here, real household credit outstanding) is determined by the first lag of the error-correction mechanism associated with a long run relationship, by its own past first differences and by the current and past first differences of the (assumed) exogenous variables.

For a long run relationship to exist we require two properties: first, its error-correction coefficient in the ARDL equation is significant for (almost) all individuals (countries);

second, at least one of the fundamentals must have a significant coefficient in this relation. To investigate these two properties, we use the “bounds tests” proposed by Pesaran *et al.* (2001) for time series.

It is worth emphasizing that the ARDL framework does not provide a test for cointegration taking advantage of the panel dimension. We refer therefore to the procedure suggested by Kao (1999) for investigating whether the long run relationship estimated in the ARDL framework can be recognized as a cointegration relationship.

The paper is organized as follows. In section 2, we propose a brief survey of credit demand analysis and describe the data we use. Section 3 discusses the interest of ARDL models. In section 4, we recall briefly how to validate the existence of long run relationships using panel cointegration tests and “bounds tests”. Section 5 presents the empirical results. Section 6 addresses the robustness of results and in particular whether one really identifies a demand as opposed to a supply schedule. Section 7 concludes.

2 Structural Analysis of Credit Markets and data sources

We just provide here a quick overview of the literature on the equilibrium on the credit market, with exclusive focus on credit to households. A standard reference is the paper by Friedman and Kuttner (1993) who identify supply and demand of loans to the private sector in the US. The workhorse of the literature is the estimation of a credit demand relationship defined as $D_t^d = f(R_t, \varepsilon_t)$, where D_t is loan volume (*i.e.* in real terms, after applying a price deflator), R_t is the interest rates and ε_t represents the set of demand variables, including aggregate wealth, disposable income (which can be viewed as a proxy for wealth), investment and cost variables. The demand for credit is proportional to these variables. An other variable that is often considered in the credit literature is inflation, which measures additional costs and has a positive effect on credit demand. For household credit, relative house prices (*i.e.* house prices divided by the private consumption deflator) are equally introduced as a proxy for wealth. Its effect on loan demand is either direct, measuring the cost of housing (higher house prices also improve the value of the collateral pledged against new credit), or indirect because of the effect of higher consumption (associated with higher wealth) on consumption credit. Wealth effects have long been considered as a major factor behind credit cycles (Iacoviello, 2005, as well as Kiyotaki and Moore, 1997, for a modern version of Fisher’s (1933) debt deflation theory). In addition, demand is negatively affected by interest rates: $\frac{\partial D_t^d}{\partial R} < 0$.¹ Considering such an equation, one can obviously draw a parallelism with the literature on money demand.

When dealing with credit markets, such an equation is often complemented with a credit supply equation, which introduces risk factors and variables measuring banks’ profitability. The latter is written as $D_t^s = f(R_t, \varepsilon_t)$, where $\frac{\partial D_t^s}{\partial R} > 0$ (see Neven and Röller, 1999, among others);

¹Another issue is whether one should consider real or nominal interest rates, but the debt burden depends on nominal interest rates.

Such an approach is also used by Fase (1995) for the Netherlands, although the author focuses on credit demand, assuming that the stock of credit is mainly determined by the demand side of the market. De Greef and de Haas (2000) for the Netherlands and Gimeno and Martinez-Carrascal (2006) for Spain include disposable income and real house prices in the demand for credit and consider the joint dynamics of house prices and mortgage loans. Hofmann (2001) investigates such a system for a set of OECD countries. It is beyond the scope of the current paper to estimate such a joint system, in particular because many other variables affect house prices, notably through demographic developments, which are actually quite exogenous to credit developments.

We focus on credit demand and look for a robust estimation method of a single equation, as in Calza *et al.* (2003), but instead of looking directly at euro area aggregates like these authors, we take advantage of the panel dimension.

Indeed, the identification of a demand equation from data is problematic because demand and supply factors operate jointly and implicitly influence the series which are observed. Even if one finds a negative estimate of the elasticity of credit to interest rate, one cannot strictly conclude that one has estimated a credit demand equation. In that case, one can just claim that demand factors are working, while the equation might just be a reduced form including both supply and demand effects. One way forward is therefore to test whether the residuals of the regression of credit onto different demand-type fundamentals are not correlated with one (or several) proper supply factor(s). If it is the case, one can conclude that the regression really identifies a demand function.

Accordingly, we first estimate a credit function by estimating a long run equation and validate the existence of cointegration between the credit variable and demand-type fundamentals. Then, we look at the residuals of the regression and check that they are not correlated with indicators of bank profitability.

The database is described with further details in Appendix A.1. It is made of quarterly data on loans to households, private consumption deflator, investment, long term interest rates and national house price indexes. The source of data are Eurosystem quarterly monetary statistics for credit and the OECD Economic Outlook database for the macroeconomic series. House prices come from a database constructed by Banque de France assembling homogenous data on price of existing houses. Due to data availability, the database only covers the 1991:1-2005:4 period and includes $N = 9$ euro area countries (Portugal, Greece and Luxembourg are always excluded), but we also exclude Belgium and Austria, for some part of the analysis, since these countries exhibit non significant results. As a consequence, $N = 7$ or 9 series over $T = 60$ periods. All series are seasonally adjusted, when necessary (this is only required for the credit series).

In what follows, we recall the main results obtained from panel data analysis to estimate long run relationships, by focusing first on the ARDL framework.

3 ARDL Models (Pesaran *et al.*, 1999)

ARDL models are widely used in the literature (see in particular Hendry *et al.*, 1984), and its reparametrisation as an ECM is well established in the time series context (Bewley, 1979, Bardsen, 1989).² The main interest of ARDL models is threefold: (i) they provide a convenient way to deal with long run relationships by focusing on the dynamics of one single equation, where the long run relationship and the short run dynamic are estimated jointly; (ii) they can therefore be easily extended to a panel framework; (iii) they allow to deal with variables that are possibly of different order of integration, namely $I(0)$ and $I(1)$, and not simply $I(1)$.

We focus on ARDL models for times series before looking at these models for panel data.

3.1 ARDL models for time series

In the ARDL approach, by Pesaran and Shin (1999), one concentrates on one endogenous variable of interest, Y , or equivalently on one single equation:

$$Y_t = a_0 + a_1 t + \sum_{j=1}^p \lambda_j Y_{t-j} + \sum_{j=0}^q \delta'_j X_{t-j} + u_t, \quad (1)$$

where X denotes the set of regressors, which are supposed to be non correlated with the residuals u .

One often finds the equivalent specification:³

$$\Delta Y_t = \phi Y_{t-1} + \beta' X_t + \sum_{j=1}^{p-1} \lambda_j^* \Delta Y_{t-j} + \sum_{j=0}^{q-1} \delta_j^{*'} \Delta X_{t-j} + a_0 + a_1 t + u_t. \quad (2)$$

The objective of this subsection is to recall the intuition behind the previous equation, and in particular, its link with VAR models and VECMs.

Starting from a VAR model for the $Z_t = (Y_t, X_t)'$ vector, the system of equations is partitioned in order to get the single equation above. More precisely, as a start, one writes the (canonical) VAR model:

$$\begin{aligned} \Phi(L)(Z_t - \mu - \gamma t) &= \varepsilon_t \\ \Leftrightarrow \Delta Z_t &= \alpha_0 + \alpha_1 t + \Pi Z_{t-1} + \sum_{i=1}^{p-1} \Gamma_i \Delta Z_{t-i} + \varepsilon_t, \end{aligned}$$

Where $\alpha_0 = -\Pi\mu + (\Gamma + \Pi)\gamma$; $\Gamma = Id - \sum_{i=1}^{p-1} \Gamma_i$; $\Pi = -(Id - \sum_{i=1}^p \Phi_i)$ and $\alpha_1 = -\Pi\gamma$. $\Omega = \begin{bmatrix} \omega_{yy} & \omega_{yx} \\ \omega_{xy} & \Omega_{xx} \end{bmatrix}$ is the variance-covariance matrix of ε_t .

²See Banerjee *et al.* (1993).

³Here, the equivalence with equation (1) requires that $\phi = -(1 - \sum_{j=1}^p \lambda_j)$ and $\beta = \sum_{j=0}^q \delta'_j$; $\lambda_j^* = -\sum_{m=j+1}^p \lambda_m$ for $j = 1, 2, \dots, p-1$; $\delta_j^{*'} = -\sum_{m=j+1}^q \delta_m$ for $j = 1, 2, \dots, q-1$.

Thus, by separating the equation of Y from the ones of the other components X , with the corresponding partition of the different matrices, one can write the ΔY -equation under the ECM type form:

$$\Delta Y_t = c_0 + c_1 t + \pi_{yy} Y_{t-1} + \pi_{yx} X_{t-1} + \sum_{i=1}^{p-1} \Gamma_{yi} \Delta Z_{t-i} + \varepsilon_{yt},$$

with Π written as $\Pi = \begin{bmatrix} \pi_{yy} & \pi_{yx} \\ \pi_{xy} & \Pi_{xx} \end{bmatrix}$ and $\varepsilon_t = (\varepsilon_{yt}, \varepsilon'_{xt})'$.

Finally, in order to orthogonalize the ΔY and ΔX innovations, one introduces contemporaneous regressors X in the ΔY -equation as following:⁴

$$\Delta Y_t = c_0 + c_1 t + \pi_{yy} Y_{t-1} + \pi_{yx.x} X_{t-1} + \sum_{i=1}^{p-1} \Psi'_i \Delta Z_{t-i} + \omega' \Delta X_t + u_t,$$

where $\pi_{yx.x} = \pi_{yx} - \omega' \Pi_{xx}$ (matrix $1 \times k$); $\omega = \Omega_{xx}^{-1} \omega_{xy}$ and $u_t = \varepsilon_{yt} - \omega_{yx} \Omega_{xx}^{-1} \varepsilon_{xt}$. With u_t *i.i.d.* $N(0, \omega_{uu})$ and $\omega_{uu} = \omega_{yy} - \omega_{yx} \Omega_{xx}^{-1} \omega_{xy}$.

If $\phi = \pi_{yy}$ and $\beta = \pi_{yx.x}$, after redefining the lag polynomial in Z in order to get the contemporaneous value of X in the level part, this yields Pesaran's *et al.* (2001) single equation (2) of the ARDL approach:

$$\Delta Y_t = c_0 + c_1 t + \pi_{yy} Y_{t-1} + \pi_{yx.x} X_t + \sum_{i=1}^{p-1} \tilde{\Psi}'_i \Delta Z_{t-i} + \omega' \Delta X_t + u_t. \quad (3)$$

By construction, innovations u_t and ε_{xt} (the canonical innovations of the X variables in the canonical VAR model) are not correlated.

Pesaran *et al.* (2001) assume that $\{\pi_{xy} = 0\}$. Such an assumption is equivalent to exclude the feedback of the level of Y on the level of X , and to assume that there exists at most one long run relationship with Y as endogenous variable. One can refer to the example provided by Pesaran, Shin and Smith (2001) who study a wage equation, where real wage is a function of labor productivity, but the effect of the level of the real wage on productivity is excluded, consistently with bargain theory. One can also test for that constraint. Here we assume that house prices are exogenous and that credit does not cause house prices. Such an assumption is consistent with Gouteron and Szpiro (2005), who indicate that excess credit does not explain house prices in the euro area, in opposition with the earlier conclusion by Hofmann (2001), but for aggregate private sector credit in the latter case.

It is worth emphasizing that the ARDL models have been introduced to avoid pre-testing to insure that all components of Z are $I(1)$ as required by the VECM specification.

⁴It is equivalent to multiply both members of the VAR specification by the matrix:

$$P = \begin{bmatrix} 1 & -\omega_{yx} \Omega_{xx}^{-1} \\ 0 & Id \end{bmatrix}.$$

If $\phi \neq 0$, Π has reduced rank $r + 1$ ($r \leq k$, the number of variables in Z) and one can write a long run relation with Y_t as endogenous variable as:

$$Y_t = \theta_0 + \theta_1 t - \frac{1}{\phi} \beta' X_t + v_t. \quad (4)$$

The long run relationship is non degenerated if $\theta = -\frac{\beta}{\phi}$, the vector of long run parameters “conditional on” X , is non null (or equivalently, if $\beta \neq 0$).

If X_t , Y_t are $I(1)$ and v_t is $I(0)$, one can claim that Y_t and X_t cointegrate according to the conditional relationship. Thus ϕ characterizes the intensity of the “error-correction” mechanism.⁵

As already indicated, the coefficients associated with X_t , namely $\frac{\pi_{yx.x}}{\pi_{yy}}$, are not identical to the ones in the canonical VECM (*i.e.* $\frac{\pi_{yx}}{\pi_{yy}}$).

3.2 Extension to panel data

When the ARDL specification is used for panel data, a single equation is written for each individual i :

$$\Delta Y_{it} = \phi_i Y_{it-1} + \beta_i' X_{it} + \sum_{j=1}^{p-1} \lambda_{ij}^* \Delta Y_{it-j} + \sum_{j=0}^{q-1} \delta_{ij}^{*'} \Delta X_{it-j} + c_{0i} + c_{1i} t + u_{it}, \quad (5)$$

$\forall 1 \leq i \leq N$ and $1 \leq t \leq T$.

The main assumptions required by of the Pesaran *et al.*'s (1999) Pooled Mean Group model are as follows:

- (a₁) Residuals u_{it} are assumed to be independent across individuals and independent from regressors X_{it} (the latter hypothesis is just necessary to get consistent estimates of the short run parameters λ_{ij}^* and $\delta_{ij}^{*'}$), but they may have different variances $\sigma_i^2 = \text{Var}(u_{it})$.
- (a₂) For each individual i , the long run relation:

$$Y_{it} = -\frac{\beta_i'}{\phi_i} X_{it} + v_{it},$$

corresponds to a cointegration relation, that is v_{it} is $I(0)$.

- (a₃) The long run coefficients $\theta_i = -\frac{\beta_i}{\phi_i}$ are the same for the different individuals under the long run homogeneity hypothesis:

$$\forall i, \theta_i = -\frac{\beta_i}{\phi_i} = \theta$$

⁵ Y_t is $I(0)$, like v_t , if $\theta = 0$, that is if $\pi_{yx} - \omega' \Pi_{xx} = 0$ with $\omega \neq 0$.

(the short run parameters can differ from an individual to another).⁶

Assumption (a_1) is often supposed to be satisfied without being tested. We examine, in the next section, how to test this hypothesis, by using the cointegration test proposed by Kao (1999) in a panel context, or the “bounds tests” procedure proposed by Pesaran *et al.* (2001) for time series analysis.

Assumption (a_3) is usually tested, by referring to an Hausman statistic measuring a distance between the estimator of the unconstrained model named Mean Group Estimator (where the long run parameters are free like the short run ones) and the estimator of constrained model named PMG estimator (under the long run homogeneity hypothesis). When there is no rejection, one concludes to long run homogeneity.

Now, the question is how to test these different assumptions and, among them, the cointegration hypothesis.

4 Investigation of long run relationships from panel data

In this section, we examine different approaches to validate the existence of a long run relationship for panel data. We first recall briefly the methodology proposed by Kao (1999) to test for cointegration by using a pooled procedure before examining how to exploit “bounds tests” on individuals countries (Pesaran *et al.*, 2001).

4.1 Cointegration tests in the lines of Kao (1999)

In the Kao (1999) framework, one assumes that the long run parameters β are the same for all individuals and allows heterogeneity through fixed effects.⁷ More precisely, the set-up is the following:

$$\begin{aligned} Y_{it} &= Z_t' \gamma_i + X_{it}' \beta + e_{it}, \\ Z_t' \gamma_i &= (\alpha_i + \delta_i t + \dots), \\ X_{it} &= X_{it-1} + \varepsilon_{it}, \end{aligned}$$

where ε_{it} is *i.i.d.*; accordingly, the variables (Y_{it}, X_{it}) are supposed to be independent for different individuals. In what follows, the deterministic part is supposed to be reduced to a constant ($Z_t = 1$).

Kao (1999) considers the estimation of β with the least squared dummy variable (LSDV) estimator:

$$\hat{\beta} = \left[\sum_{i=1}^N \sum_{t=1}^T \tilde{X}_{it} \tilde{X}_{it}' \right]^{-1} \left(\sum_{i=1}^N \sum_{t=1}^T \tilde{X}_{it} \tilde{Y}_{it} \right),$$

⁶The long term coefficients are estimated by maximisation of a concentrated likelihood function, through an iterative procedure (“Newton-Raphson” or “back-substitution” algorithm), introducing the θ_i , ϕ_i and σ_i^2 coefficient (pooled estimation). Using the estimated $\hat{\theta}_i$ coefficients (as derived from the previous algorithms), the short run coefficients (including the ϕ_i , σ_i^2 and the intercepts) are then estimated separately for each country by OLS. The PMG estimator for the short run coefficients is the average over all countries.

⁷This formulaion is less restrictive than a Pooled model which specifies constant coefficients.

with $\tilde{Y}_{it} = M_{\{Z\}} Y_{it}$ and $\tilde{X}_{it} = M_{\{Z\}} X_{it}$ (Frisch-Waugh).⁸

The cointegration tests, in the panel context, are thus Unit Root tests on the estimated residuals:

$$\hat{e}_{it} = \tilde{Y}_{it} - \tilde{X}'_{it} \hat{\beta}.$$

By implementing the following regressions :

$$\hat{e}_{it} = \rho \hat{e}_{it-1} + v_{it}, \quad (\text{A})$$

$$\text{and } \hat{e}_{it} = \rho \hat{e}_{it-1} + \sum_{j=1}^p \varphi_j \Delta \hat{e}_{it-j} + v_{itp}, \quad (\text{B})$$

Kao (1999) tests the null hypothesis of non cointegration, namely $H_0 : \rho = 1$, against $H_1 : \rho < 1$, and suggests from (A), four Dickey-Fuller type statistics (DF_ρ , DF_t , DF_ρ^* and DF_t^*) and from (B), an Augmented Dickey-Fuller type statistic (ADF_t). While the DF_ρ and DF_t statistics assume strong exogeneity of regressors and errors and their parameters depend on nuisance parameters, the DF_ρ^* and DF_t^* statistics take into account the possible endogeneity between regressors and errors (see Appendix C.1). By construction, DF_ρ^* , DF_t^* and ADF_t ⁹ do not depend on nuisance parameters and follow, according to a sequential asymptotic theory,¹⁰ a standard normal distribution. In what follows, we only refer to the last three statistics and we prefer to use the ADF statistic, because the associated unit-root regression can be proved to be a constrained version of the regression implemented to estimate the ARDL equation (see Appendix C.2).

Now, we turn to the “bounds tests” procedure.

4.2 “Bounds tests” in the lines of Pesaran *et al.* (2001)

For individual time series, starting from the VECM (equation (3)), Pesaran *et al.* (2001) test the error-correction and long term coefficients, *i.e.* the following null hypothesis:

$$H_0 : \{\pi_{yy} = \pi_{yx.x} = 0\},$$

against the alternative $H_1 : \{\pi_{yy} \neq 0 \text{ or } \pi_{yx.x} \neq 0\}$.

The test statistic has a Fisher distribution, which depends on the integration order of series Y and X and also on the deterministic part of the long run equation. For example, at a significance level of 5%, with no deterministic component, the critical value is equal to 5.73 if both series are $I(1)$ and 4.94 if they are both $I(0)$.¹¹

When the statistic is smaller than a lower bound, the null hypothesis is not rejected, while it is rejected when the statistic is greater than an upper bound (5.73 for example, in

⁸ $M_{\{Z\}} = \left[I_N \otimes \left(I_T - \frac{e_T e'_T}{T} \right) \right]$ which equivalent to “Within” operator W_N ; where I is identity matrix, $e'_T = (111\dots 1)$ and $Z = Z'_i \gamma_i$.

⁹ ADF_t statistic used to take into account both endogeneity of regressors and serial correlations of residuals.

¹⁰ $T \rightarrow \infty$ followed by $N \rightarrow \infty$.

¹¹ See tables in Pesaran *et al.* (2001), for more than one regressor.

the case of two $I(1)$ series). One can not conclude between the two bounds. The previous test allows to test for the existence of a relationship between the levels of the different series, whatever the stationarity properties of the regressors (TS or DS). It consists in testing whether the Long run parameters of equation (3) are jointly equal to 0.

Under the null hypothesis, the asymptotic distribution of the Fisher statistic is not standard, whatever the integration order of the regressors ($I(0)$ or $I(1)$).

Pesaran *et al.* (2001) also test for the null hypothesis:

$$H_0 : \{\pi_{yy} = 0\},$$

against the alternative $H_1 : \{\pi_{yy} \neq 0\}$.

In what follows, we refer to the tables for the distribution of the $t_{\pi_{yy}}$ statistic,¹² presented by Pesaran *et al.* (2001). Note however that in the panel context, Pesaran *et al.* (1999) conclude that π_{yy} is significant by comparing its estimate to the corresponding standard error and observing that it is “highly significant” without referring to any table.

In order to test for cointegration, one needs to implement a sequential procedure. At the first step, one tests for the joint nullity of π_{yy} and $\pi_{yx.x}$. If the null hypothesis is not rejected, one can be sure that Y and X do not cointegrate according to ΔY -equation. If the null is rejected, one has to look at π_{yy} and $\pi_{yx.x}$ successively and test whether they are significant. At the second step, one implements the previous test for $\pi_{yy} = 0$. If nullity of π_{yy} is not rejected, one has to exclude any long run relationship including Y . If nullity of π_{yy} is rejected, one should also test, in a third step, for the presence of X in the cointegration relation, *i.e.* one has to explicitly test for $H_0 : \{\pi_{yx.x} = 0\}$.

Unfortunately, this is not available in Pesaran *et al.*'s (2001) approach and would require an extension of their testing strategy. However, if one knows a priori that Y is $I(1)$, then one can conclude that Y and X cointegrate and that $\pi_{yx.x} \neq 0$, once $\pi_{yy} \neq 0$. Indeed, the joint conditions $\pi_{yx.x} = 0$ and $\pi_{yy} \neq 0$ would imply that Y is $I(0)$ like the right member of the ARDL single equation rewritten as:

$$-(\pi_{yy}Y_{t-1} + \pi_{yx.x}X_{t-1}) = c_0 + c_1t + \Delta Y_t + \sum_{i=0}^{p-1} \Psi'_i \Delta Z_{t-i} + \omega' \Delta X_t + u_t.$$

The extension of the results obtained by Pesaran *et al.* (2001) is beyond the topic of the present paper, and we only consider the first two steps.

5 Empirical results

We consider that credit demand depends on long run interest rates, as well as two scale variables, namely households' investment and house prices. Credit is expressed in real terms :

$$\ln(D_{it}/P_{it}) = \alpha_i + \beta_i t + \gamma_i \ln(INV_{it}) + \tau_i LTR_{it} + \zeta_i \ln(PLOG_{it}) + e_{it} \quad (6)$$

¹²In principle, one should also have verified before that ΔY does not cause ΔX in Granger's sense. We assume that it is the case in our empirical study.

where D_{it} is the stock of credit to households in country i at date t , INV_{it} is the investment variable, LTR_{it} is the long term nominal interest rate, β_i could measure the financial development trend, *i.e.* the tendency of financial assets/liabilities to grow more rapidly than GDP or income.¹³ Note that $\tau < 0$ is consistent with a demand equation and γ_i is expected to be positive and close to 1. The empirical results are not satisfactory for specifications including the short run interest rate, which is not the usual reference for pricing loans to households. Indeed, housing loans represent usually more than 80% of total credit to households. $PLOG$ is the real house price (*i.e.* divided by the consumption deflator).

The objective of the empirical strategy is to estimate such an equation including the long run equilibrium and the short run dynamics. Estimating directly the previous equation, independently across countries (see “regression 1” below), often leads to the rejection of cointegration, due to the lack of power of usual tests. Taking advantage of the panel dimension we first test for cointegration, then implement the Pesaran *et al.*'s (1999) PMG approach.

To summarize, the different models that we estimate are the following :

Regression 1: Unrestricted country-by-country equation (whose average of coefficients yields the Mean Group or “MG” estimator), where Y_{it} stands for real credit, *i.e.* $\ln(D_{it}/P_{it})$, and X_{it} for the different explanatory variables:

$$\Delta Y_{it} = \phi_i (Y_{it-1} - \theta'_i X_{it}) + \sum_{j=1}^{p-1} \lambda_{ij}^* \Delta Y_{it-j} + \sum_{j=0}^{q-1} \delta_{ij}^{*'} \Delta X_{it-j} + \mu_i + \varepsilon_{it} \quad (\text{Regr. 1})$$

$\forall 1 \leq i \leq N$ and $1 \leq t \leq T$.

Regression 2 : ARDL-ECM with common long run and free short run coefficients, namely the Pooled Mean Group or “PMG” estimator:

$$\Delta Y_{it} = \phi_i (Y_{it-1} - \theta' X_{it}) + \sum_{j=1}^{p-1} \lambda_{ij}^* \Delta Y_{it-j} + \sum_{j=0}^{q-1} \delta_{ij}^{*'} \Delta X_{it-j} + \mu_i + \varepsilon_{it} \quad (\text{Regr. 2})$$

$\forall 1 \leq i \leq N$ and $1 \leq t \leq T$.

Regression 3 : Dynamic fixed effects (DFE), which assumes short and long run homogeneity (except the constant term):

$$\Delta Y_{it} = \phi (Y_{it-1} - \theta' X_{it}) + \sum_{j=1}^{p-1} \lambda_j^* \Delta Y_{it-j} + \sum_{j=0}^{q-1} \delta_j^{*'} \Delta X_{it-j} + \mu_i + \varepsilon_{it} \quad (\text{Regr. 3})$$

$\forall 1 \leq i \leq N$ and $1 \leq t \leq T$.

Regression 2 is the model that we estimate, using the Schwarz criterion to validate the common lag structure. Starting from the unconstrained regression 1, we use a Hausman test to assess whether the homogeneity constraints can be accepted.

¹³None of the empirical results includes a deterministic trend, as it turns out not to be significant over the sample period.

5.1 Results from ARDL

We consider the model with households' investment, as well as the house price index in the long run relation as well as in the short run dynamics.

We first proceed with the Pooled Mean Group (PMG) approach for all countries. Table 1 provides the long run coefficients for the three regressions. In the PMG model, *i.e.* when constraining the 9 countries to have the same long run relationship on the level of the variables as in regression 2, the long run coefficients have the right sign. The long run elasticity of credit to investment is around 1.6 and the elasticity of house prices is 0.57, while the semi-elasticity of long term interest rates is -0.1 . The Hausman test for overall homogeneity is not rejected indicating that PMG regression is supported by the data, in the sense that regression 2 is not statistically different from the average of the long run coefficients from regression 1, exhibited in column.

**Table 1: Panel estimates (9 countries)
model with invest., long term nominal int. rates and house prices**

lag 1/1/0/1	PMGE		MGE		DFE		H-test ⁽¹⁾
	coef	t-ratio	coef	t-ratio	coef	t-ratio	p-value
$\ln(INV)$	1.578	7.279	1.329	1.128	2.642	2.124	0.83 (indiv.)
LTR	-0.098	-10.923	-0.053	-0.828	-0.158	-1.942	0.47 (indiv.)
$\ln(PLOG)$	0.570	6.611	0.846	1.193	0.151	0.356	0.70 (indiv.)
ϕ	-0.030	-4.779	-0.037	-1.965	-0.016	-1.621	0.74 (joint)

⁽¹⁾Hausman test comparing PMGE and MG results

PMGE : Pooled Mean Group Est.; MGE : Mean Group Est.; DFE : Dynamic Fixed Effect

However, when looking at the error-correction coefficient ϕ_i in the different countries (Table 2), it appears that it is not significant for two countries, namely Austria and Belgium. For the other countries, the equations exhibit good properties, in particular there is no autocorrelation of residuals. In the case of Italy, it is accepted but at the 1% level.

It is worth noting that assumption (a_1) of the Pesaran *et al.*'s (1999), PMG estimator is satisfied since we verify that there is no correlation of the residuals across countries according to a Pearson's test (see Appendix A.4 for results).¹⁴

Regarding the speed of adjustment as provided by the error-correction adjustment ϕ_i , lower adjustment is observed in Italy and Ireland, while higher adjustment appears in the Netherlands and Germany.

¹⁴The statistic of the Pearson test of significance of the correlation coefficient r is $\frac{r\sqrt{n-2}}{\sqrt{1-r^2}}$ follows a Student distribution with $(n-2)$ degrees of freedom, where n is the number of observations.

Table 2: ECM coefficients in PMGE (9 countries)
model. with invest., long term nominal int. rates and house prices

country	ϕ_i	std-error	t-ratio	Resid autocorr. test (\dagger)	\bar{R}^2
Austria	0.003	0.006	0.460	0.02	-0.06
Belgium	-0.010	0.008	-1.272	4.17	-0.08
Germany	-0.048 (***)	0.006	-7.644	1.20	0.76
Spain	-0.030 (***)	0.007	-4.474	0.02	0.59
Finland	-0.038 (***)	0.005	-7.166	1.13	0.56
France	-0.038 (***)	0.005	-7.470	0.53	0.70
Ireland	-0.025 (***)	0.005	-4.790	0.61	0.51
Italy	-0.026 (***)	0.008	-3.222	5.84	0.25
Netherlands	-0.061 (***)	0.015	-4.178	0.87	0.25

PMGE : Pooled Mean Group Estimate; (\dagger) Godfrey's test distributed as $\chi^2(1)$.
(***) Significance at 1% according to Student distribution.

The heterogeneity of short run dynamics also appears from comparing the estimates of the long run elasticities obtained for the Dynamic Fixed effects specification (DFE, 3rd column in Table 1) and the ones obtained with the PMG specification. As the introduction of different constraints on the short run dynamics (compare regr.2 and regr.3) does seem to matter, it implies that the short run dynamics is not identical across all countries. So we decide to drop Austria and Belgium and proceed again to the PMG approach for 7 countries. The results are almost unchanged for the PMG estimation (see Table 3), but the DFE estimation is now very close to the PMG.¹⁵

Table 3: Panel estimates (7 countries)
model with invest., long term nominal int. rates and house prices

lag 1/1/0/1	PMG		MG		DFE		H-test ⁽¹⁾
	coef	t-ratio	coef	t-ratio	coef	t-ratio	p-value
$\ln(INV)$	1.593	7.257	1.834	1.246	1.517	6.885	0.87 (indiv.)
LTR	-0.098	-10.848	-0.042	-0.524	-0.100	-5.663	0.49 (indiv.)
$\ln(PLOG)$	0.565	6.508	0.895	0.965	0.533	4.058	0.72 (indiv.)
ϕ	-0.038	-7.809	-0.033	-2.171	-0.033	-5.392	0.24 (joint)

⁽¹⁾Hausman test comparing PMGE and MG results

PMGE : Pooled Mean Group Est.; MGE : Mean Group Est.; DFE : Dynamic Fixed Effect

Furthermore, the results appear stable over time on the basis of recursive CUSUM tests applied to the constrained PMG model with common long run relationship, or the unconstrained model,¹⁶ that we discuss more fully in the following section. The stability over time of country models exhibiting different short run dynamics leads to the conclusion

¹⁵Indeed, the DFE estimation is also supported by the data on the basis of a Hausman test, comparing MG and DFE, but we rely in the rest of the paper on the PMG model which is the most general specification. In addition, the existence of differences in the adjustment coefficient also favours the PMG model.

¹⁶Results for CUSUM tests are available upon request from the authors.

that countries share common long run dynamics (as verified by the PMG specification), while institutional idiosyncrasies are still at play and persistent in the short run.

Such a conclusion, regarding the relatively good fit of the PMG model, is also apparent from simulating the models. The charts below exhibit the static as well as the dynamic simulation. In the first case, the model of regression 2 uses historical values for the lagged value of the endogenous variable, while in the latter case the estimated value is used recursively. The dynamic simulation is obviously a more stringent test of the model fit. It turns out that the dynamic simulation follows quite closely the actual year-on-year growth of real credit to households. This is more significantly so in the case in Germany and France, as well as Spain and the Netherlands. On the other hand, large discrepancies appear notably for Italy, but also for Finland and Ireland.

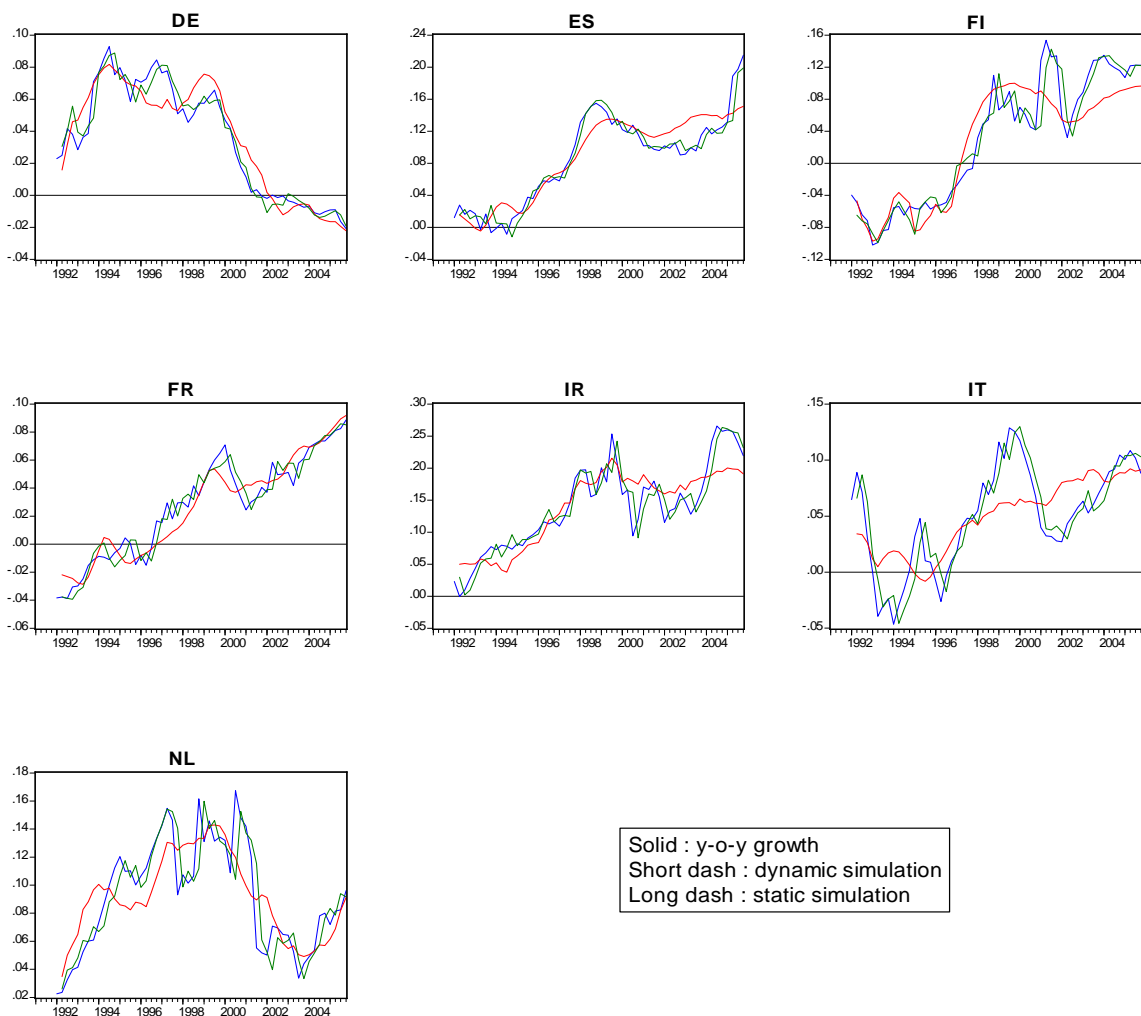


Figure 1 : Static and dynamic simulation of y-o-y real credit growth

The conclusion at this stage is that credit behaviour exhibit similar long run behaviour across a significant number of countries in the euro area (France, Germany, Spain and to a smaller extent in Italy), while Ireland and Finland also share some common features, and the other countries (Austria and Belgium) still follow specific dynamics. This provides evidence of converging financial behaviour, even if short run dynamics, linked to institutions, are still different.

6 Robustness Analysis

We consider now two types of robustness analysis. First, we investigate whether there is indeed cointegration and whether we have identified a demand equation.

6.1 Cointegration Analysis

As indicated before, Pesaran's PMG model rests on the assumption that the variables are cointegrated and we proceed now to verify that assumption. Anticipating on the results of the subsection, we show that looking at various methods, either based on the panel structure of the data, or on the time series dimension, it appears that cointegration is only accepted for 5 to 6 countries out of 7, Italy exhibiting a different behaviour than the other countries.

6.1.1 Kao's panel approach

On the basis of Kao's approach, it appears that cointegration is rejected for our set of 9 countries and also on the set of 7 countries. To get evidence of cointegration, if one refers to the ADF statistic, which is the statistic we prefer for the reasons explained before, one needs to reduce the sample to 6 countries (p-value for ADF_t is 10.6%) or to 5 countries (p-value for ADF_t is 0,09%). In the following table, we report the values of the results obtained with the different test statistics for different sets of countries.

**Table 4: Kao (1999) cointegration test
(Model with invest. and L.T. nominal interest rates and house prices)**

Stat. of tests		DF_{ρ}^*		DF_t^*		ADF_t	
Subgroup	Countries	t-stat	p-val	t-stat	p-val	t-stat	p-val
7 countries	DE,ES,FI,FR,IR,IT,NL	-0.254	0.399	0.222	0.588	-0.771	0.220
6 countries	DE,ES,FI,FR,IT,NL	-0.853	0.196	-0.266	0.395	-1.246	0.106
5 countries	DE,ES,FR,IT,NL	-1.217	0.112	-0.312	0.377	-1.334	0.091
4 countries	DE,ES,FR,NL	-2.649	0.004	-0.892	0.186	-1.862	0.031
4 countries	DE,ES,IT,NL	-1.809	0.035	-0.717	0.237	-1.519	0.064
3 countries	DE,ES,NL	-4.016	0.000	-1.504	0.066	-2.213	0.013

6.1.2 Time series approach to cointegration

Further evidence in favour of cointegration on a limited set of countries can be found from country-by-country analysis, either testing the significance of the error-correction mechanism or using the more restrictive “bounds tests”.

Unconstrained models We first study the traditional approach of cointegration,¹⁷ examining the significance of the ECM coefficient in the individual time series. We study therefore the properties of unconstrained models, where the long run relationship is country-specific, according to the traditional time series approach. It turns out that Finland and Ireland have no longer a significant adjustment mechanism (see Table 5).

More generally, one concludes that the intensity of error-correction mechanism - measured by the ϕ_i coefficient - varies with the country. This result is in favor of short term heterogeneity.

**Table 5: ECM coefficients country-by-country (unconstrained)
Model. with invest., long term nominal int. rates and house prices**

Country	ϕ_i	std-err.	Lag struc.	inf. crit. ⁽²⁾	Autocorr. (†)	\bar{R}^2
DE	-0.058 (***)	0.011	1/1/0/2	<i>SBC</i>	0.13	0.80
ES	-0.111 (***)	0.034	3/3/2/3	<i>AIC</i>	0.03	0.79
FI	-0.015	-0.016	3/1/0/0	<i>AIC</i>	0.01	0.64
FR ⁽¹⁾	-0.541 (**)	0.119	1/3/3/2	<i>SBC</i>	5.92	0.82
IR	0.012	0.024	1/1/0/1	<i>AIC/HQ</i>	0.00	0.54
IT	-0.099 (***)	0.041	1/0/3/1	<i>AIC/HQ/SBC</i>	0.10	0.47
NL ⁽¹⁾	-0.212 (***)	0.059	1/2/0/0	<i>SBC</i>	0.97	0.34

⁽¹⁾Models with unrestricted intercept and unrestricted trend

⁽²⁾*AIC*, *SBC* and *HQ* are resp. Akaike, Schwarz Bayesian and Hannan-Quinn information criteria

Significance level : (***) 1%, (**) 5% and (*) 1%; (†) Godfrey’s test distributed as $\chi^2(1)$.

“Bounds tests” Using the “bounds tests” advocated by Pesaran *et al.* (2001), we validate the joint significance of the long run and the adjustment coefficients (column 1 of Table 6, based on a Fisher test). However, it appears to be more difficult to validate the significance of the ϕ_i coefficient only (column 2 in Table 6), since our t-statistics are below the high critical value suggested by Pesaran *et al.* (2001). The null hypothesis of non cointegration is only rejected for Germany and France, while inconclusive results are found for Spain and the Netherlands.

It should be kept in mind that such a test is run on individual time series and do not take advantage of the panel nature of our database, since the critical values are not

¹⁷For illustration purposes, we present results implementing Banerjee *et al.*’s (1998) approach, with a simple test of significance (Student t) of the error-correction mechanism ϕ_i from an OLS estimation of the ECM model.

available in the panel context. One can think that our series are too short to display cointegration properties.

Indeed, Hofmann (2001) who uses time series similar to ours but over a longer period (1980-1998) validates cointegration for 16 OECD countries and his estimates of the long run coefficients are rather close to our estimates (with average values of the output, interest rate and property price coefficients about 1.3, -0.02 and 0.60 respectively).

Table 6 : Bounds Testing on unconstrained models

Tests	$H_0^{\pi_{yy}} : \pi_{yy} = 0, H_0^{\pi_{yx.x}} : \pi_{yx.x} = 0'$; $H_1^{\pi_{yy}} : \pi_{yy} \neq 0, H_1^{\pi_{yx.x}} : \pi_{yx.x} \neq 0'$;	$H_0^{\pi_{yy}} : \pi_{yy} = 0$ $H_1^{\pi_{yy}} : \pi_{yy} \neq 0$
Country	F-statistic	t-statistic
DE	30.84 (***)	-5.51 (**)
ES	4.40 (**)	-3.23 (inconclusive inference)
FI	8.2 (***)	-0.93
FR ⁽¹⁾	7.92 (***)	-4.52 (**)
IR	8.48 (***)	0.48
IT	6.97 (***)	-2.42
NL ⁽¹⁾	6.99 (***)	-3.60 (inconclusive inference)
F-stat : I(1) \rightarrow 5.61 (at 1% (***)) and 4.35 (at 5% (**)); I(0) \rightarrow 4.29 (at 1%) and 3.23 (at 5%)		
F-stat Unrest ⁽¹⁾ : I(1) \rightarrow 6.36 (at 1%) and 5.07 (at 5%); I(0) \rightarrow 5.17 (at 1%) and 4.01 (at 5%)		
t-stat : I(1) \rightarrow 4.37 (at 1%), -3.78 (at 5%) and -3.46 (at 10%);		
t-stat : I(0) \rightarrow -3.43 (at 1%), -2.86 (at 5%) and -2.57 (at 10%)		
t-stat Unrest ⁽¹⁾ : I(1) \rightarrow -4.73 (at 1%), -4.16 (at 5%) and -3.84 (at 10%);		
t-stat Unrest ⁽¹⁾ : I(0) \rightarrow -3.96 (at 1%), -3.41 (at 5%) and -3.13 (at 10%)		
⁽¹⁾ Models with unrestricted intercept and unrestricted trend (for FR and NL)		

So, the results we obtain with the “bounds tests” do not really challenge the ones based on Kao test and presented in section 6.1.1.

Even if cointegration is not accepted for all the countries in the panel, the long run coefficients are quite stable over the sample of countries. As indicated in the table in Appendix A.2, and as compared with Table 3, the long run coefficients are relatively unaffected for the different subsamples of countries we consider. It also implies, nevertheless, that the dynamic simulations for the four “core” countries, namely Germany, France, Spain and the Netherlands are not significantly improved when the PMG long run relationship is computed on this smaller set of countries (see Appendix A.3).¹⁸

6.2 Analysis of Supply vs Demand shocks

As a further robustness check of our previous approach, we now consider whether we have truly identified a demand equation.

¹⁸In addition, the absence of cointegration for panel data has different implication than in the pure time series context, since the estimator remains consistent in the former case, while it creates spurious results in the latter case (Entorf, 1997; Kao, 1999).

Table 7: Correlation of residuals with changes in bank profits mod. w/ invest., long term int. nominal and house prices

country	gross inc./assets	net inc./assets	prof. bef. tax/assets	gross inc./loans
Germany	-0.07	0.07	0.17	-0.08
Spain	-0.08	0.00	0.09	-0.19
Finland	0.26	0.04	0.07	0.27
France	-0.16	-0.28	0.27	-0.37
Ireland	0.05	-0.18	0.10	-0.70 (***)
Italy	0.15	0.13	0.39	-0.25
Netherlands	-0.35	-0.19	-0.17	0.65

(***) Signif. diff. from zero at 1% w/ Pearson test stat. $\frac{r\sqrt{n-2}}{\sqrt{1-r^2}}$, distributed as Student t

We study whether the equation we estimated identifies credit demand, as opposed to credit supply behaviour. We test whether the residual of the equation are correlated with indicators measuring credit supply behaviour and in particular bank profitability. For that purpose, we use data collected by OECD (2005). We compute several indicators of profitability: (1) gross interest income/total assets, (2) net interest income/total assets, (3) profit before tax/total assets and (4) gross interest income/total loans. The first three indicators are overall indicators of profitability, while the latter one measures the interest margin on loans.

Data from OECD are only available at the annual frequency for the period 1991-2004 and we average the dynamic residuals from the PMG equation (*i.e.* the ε_{it} 's in regression 2). It turns out that for almost all indicators the correlation is not significantly different for zero (Table 7).

The only exception is Ireland for indicator (4). Indeed, the PMG model appears to be unsatisfactory for Ireland. We conclude therefore that, except for Ireland, we mainly capture credit demand behaviour for most countries.

7 Conclusion

In the paper, we focus on the credit demand behavior of households in the euro area countries. In particular, we examine whether a common behavior can be captured through a unique long run equation, with transitory specific features in the short run dynamics. So, we look for long term relationships within the framework of an ARDL model which allows testing for homogeneity of the long run (and the short run) dynamics. We validate the existence of a long run equation between credit volume, investment, long run nominal interest rate and an additional "fundamental", namely a relative house-price index and we interpret this equation as a demand equation, as the estimated corresponding residuals are found not to be correlated with supply factors like bank profitability. This long run equation can be constrained to be homogeneous for 7 European countries and cointegration is validated for the five larger countries. In addition, short run homogeneity is rejected, indicating that countries have only partially converged in terms of financial structures.

The question of why more convergence appears between the larger countries is reserved

for future work. One can mention, however, at this stage that the smaller countries whose banking system is more open to the rest of the world, hence more subject to external shocks (including cross country M&As, etc.), are more likely to experience possible regime shifts.

References

- [1] Banerjee, A., J. J. Dolado, J. W. Galbraith and D. F. Hendry (1993), *Co-Integration, Error-Correction and the Econometric Analysis of Non-stationary Data*, Oxford University Press, Oxford.
- [2] Banerjee, A., J. J. Dolado and R. Mestre (1998), "Error-correction mechanism tests for cointegration in a single-equation framework," *Journal of Time Series Analysis* , 19, 269-283.
- [3] Bardsen, G. (1989), "Estimation of Long Run coefficients in Error-correction Models," *Oxford Bulletin of Economics and Statistics*, 51(3), 345-450.
- [4] Bernanke, B. (1988), "Monetary policy transmission: through money or credit?," *Business Review*, Federal Reserve Bank of Philadelphia, 11, 3-11.
- [5] Bewley, R. (1979), "The Direct Estimation of the Equilibrium Response in a Linear Dynamic Model," *Economics Letters*, 3, 357-361
- [6] Calza, A., C. Gartner and J. Sousa (2003), "Modelling the Demand for Loans to the Private Sector in the Euro Area," *Applied Economics*, 35(1), 107-117.
- [7] De Greef, J. J. M. and R. T. A. de Haas (2000), "Housing Prices, Bank Lending and Monetary Policy," Research Series Supervision No 31, De Nederlandsche Bank.
- [8] European Central Bank (2002), *Report on Financial structures*, ECB, Frankfurt am Main.
- [9] Entorf, H. (1997), "Random walks with drifts: Nonsense regression and spurious fixed-effect estimation," *Journal of Econometrics*, 80, 287-296.
- [10] Fase, M. M. G. (1995), "The Demand for Commercial Bank Loans and the Lending Rate", *European Economic Review*, 39(1), 99-115.
- [11] Fisher, I. (1933), "The debt-deflation theory of great depressions," *Econometrica*, 1, 337-357.
- [12] Friedman B. M. and K. N. Kuttner (1993), "Another look at the evidence on money-income causality," *Journal of Econometrics*, 57(1-3), 189-203.
- [13] Gimeno, R. and C. Martinez-Carrascal (2006), "The Interaction Between House Prices and Loans for House Purchase: the Spanish Case," Banco de Espana, *Working Paper* 0605.

- [14] Gouteron, S. and D. Szpiro (2005), “Excess liquidity and asset prices,” Banque de France, *Notes d’Etudes et de Recherche*, # 131
- [15] Hendry, D. F., A. R. Pagan and J. D. Sargan (1984), “Dynamic specification”. In Griliches, Z. and M. D. Intriligator (ed) *Handbook of Econometrics*, vol. 2-3, Chapter 18, Amsterdam, North Holland.
- [16] Hofmann, B. (2001), “The determinants of private sector credit in industrialised countries : do property prices matter?,” Bank for International Settlements, *Working Paper* # 108.
- [17] Iacoviello, M. (2005), “House Prices, Borrowing Constraints, and Monetary Policy in the Business Cycle”, *American Economic Review*, 95(3), 739-764.
- [18] Kao, C. D. (1999), “Spurious Regression and Residual-Based Tests for Cointegration in Panel-data,” *Journal of Econometrics*, 90(1), 1-44.
- [19] Kiyotaki, N. and J. Moore (1997), “Credit Cycles,” *Journal of Political Economy*, 105(2), 211-48.
- [20] Neven, D., L. H. Röller (1999), “An aggregate Structural Model of Competition in the European Banking Industry,” *International Journal of Industrial Organization*, 17(7), 1059-1074.
- [21] Pesaran, M. H, Y. Shin and R. P. Smith (1999), “Pooled mean group estimation of dynamic heterogeneous panels,” *Journal of the American Statistical Association*, 94, 621-634.
- [22] Pesaran, M. H., Y. Shin and R. P. Smith (2001), “Bounds Testing Approaches to the Analysis of Level Relationships,” *Journal of Applied Econometrics*, 16, 289-326.
- [23] Pesaran, M. H., and Y. Shin (1999), “An Autoregressive Distributed Lag Modelling Approach to Cointegration Analysis” in “Econometrics and Economic Theory in the 20th Century”, S. Strom (ed.), chapter 11, Cambridge University Press.

A Appendix

A.1 Data description

Data for household credit demand come from Eurosystem Statistics on the period 1991:1-2005:4, backdated with national data available from BIS. The other macroeconomic indicators are from OECD Economic Outlook database. Data on house prices come from a variety of national prices on existing dwellings, but in Italy where the data cover new houses. In addition, in the case of Germany, annual figures have been interpolated to yield quarterly figures.

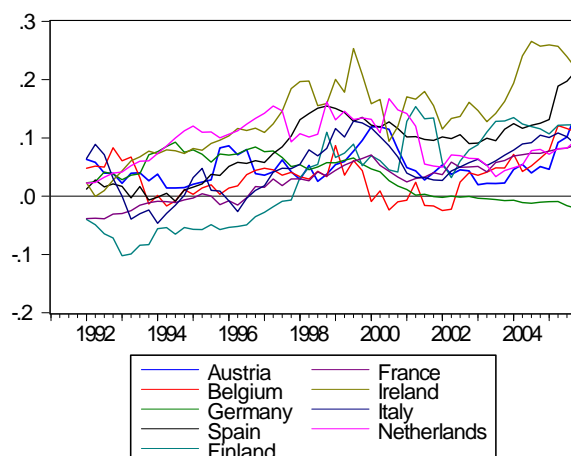


Figure 2 : Household credit demand (y-o-y growth)

A.2 Pooled Mean Group estimates on different samples of countries

Table A: Estimation of cointegration relationship
Model. with invest.long term nominal int. rates and house prices

Subgroup	$\ln(INV)$			LTR			$\ln(PLOG)$		
	PMGE	MGE	DFE	PMGE	MGE	DFE	PMGE	MG	DFE
6 countries	1.560	2.622	1.777	-0.098	-0.103	-0.093	0.571	0.532	0.522
DE,ES,FI,FR,IT,NL	(7.150)	(1.783)	(5.636)	(-10.894)	(-1.631)	(-6.032)	(6.614)	(0.527)	(4.041)
5 countries	1.533	3.077	1.507	-0.095	-0.043	-0.087	0.549	-0.262	0.546
DE,ES,FR,IT,NL	(7.033)	(1.796)	(6.447)	(-10.480)	(-1.755)	(-7.630)	(6.314)	(-0.342)	(4.024)
4 countries	1.355	1.295	1.554	-0.089	-0.072	-0.081	0.653	0.672	0.564
DE,ES,FR,NL	(7.548)	(11.673)	(4.959)	(-10.425)	(-9.010)	(-11.171)	(8.779)	(4.051)	(4.248)
4 countries	1.348	3.489	1.250	-0.092	-0.037	-0.083	0.723	-0.392	0.686
DE,ES,IT,NL	(6.400)	(1.626)	(18.853)	(-9.825)	(-1.210)	(-11.198)	(8.200)	(-0.403)	(8.873)
3 countries	1.266	1.198	1.351	-0.083	-0.069	-0.076	0.778	0.823	0.697
DE,ES,NL	(0.902)	(15.820)	(36.078)	(-8.828)	(-6.621)	(-44.762)	(10.443)	(8.450)	(7.936)

A.3 Dynamic simulation of the PMG models (regression 2 in section 5) with long run equilibrium defined on 4 countries only

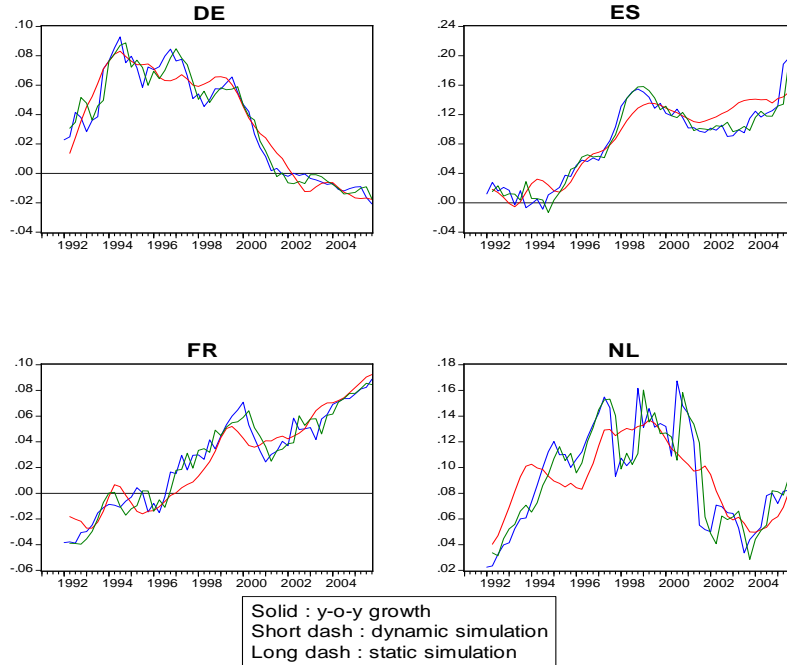


Figure 3: Dynamic simulation on the PMG model, when the long run is estimated on 4 countries

A.4 Correlation between residuals in the dynamic equation (regression 2 in section 5)

Table B : Cross-correlation of dynamic resid. model. w/ invt., long.term nominal int. rates and house prices

Country	DE	ES	FI	FR	IR	IT	NL
DE	.	-0.22(**)	-0.03(**)	0.03(**)	0.13(**)	0.16(**)	-0.03(**)
ES	.	.	-0.10(**)	0.05(**)	-0.09(**)	0.26	0.16(**)
FI	.	.	.	0.07(**)	0.23(**)	-0.15(**)	0.10(**)
FR	0.17(**)	0.21(**)	-0.01(**)
IR	0.09(**)	-0.24(**)
IT	-0.02(**)
NL

Pearson test of significance of the correlation coefficient r
 (**): significantly uncorrelated according Pearson test at 5%

B The ARDL specification

B.1 VECM characterization of the dynamics:

The VECM is written as:

$$\begin{cases} \Delta Y_t = c_0 + c_1 t + \pi_{yy} Y_{t-1} + \pi_{yx} X_{t-1} + \sum_{i=1}^{p-1} \Gamma_{yi} \Delta Z_{t-i} + \varepsilon_{yt} \\ \Delta X_t = a_0 + a_1 t + \pi_{xy} Y_{t-1} + \Pi_{xx} X_{t-1} + \sum_{i=1}^{p-1} \Gamma_{xi} \Delta Z_{t-i} + \varepsilon_{xt} \end{cases}$$

with $\Pi = \begin{bmatrix} \pi_{yy} & \pi_{yx} \\ \pi_{xy} & \Pi_{xx} \end{bmatrix}$. If one multiplies both members of the previous equation by the matrix:

$$P = \begin{bmatrix} 1 & -\omega_{yx} \Omega_{xx}^{-1} \\ 0 & Id \end{bmatrix}.$$

Thus, the ΔY and ΔX innovations are orthogonalized and one gets the equivalent characterization:

$$\begin{cases} \Delta Y_t = c_0 + c_1 t + \pi_{yy} Y_{t-1} + \pi_{yx.x} X_{t-1} + \sum_{i=0}^{p-1} \Psi'_i \Delta Z_{t-i} + \omega' \Delta X_t + u_t \\ \Delta X_t = a_0 + a_1 t + \pi_{xy} Y_{t-1} + \Pi_{xx} X_{t-1} + \sum_{i=1}^{p-1} \Gamma_{xi} \Delta Z_{t-i} + \varepsilon_{xt} \end{cases}$$

$$\Leftrightarrow \begin{cases} (1 - (1 + \pi_{yy})L)Y_t = c_0 + c_1 t + \pi_{yx.x} X_{t-1} + \sum_{i=0}^{p-1} \Psi'_i \Delta Z_{t-i} + \omega' \Delta X_t + u_t \\ (Id - (Id + \Pi_{xx})L)X_t = a_0 + a_1 t + \pi_{xy} Y_{t-1} + \sum_{i=1}^{p-1} \Gamma_{xi} \Delta Z_{t-i} + \varepsilon_{xt} \end{cases}$$

$c_0 = \alpha_{y0} - \omega' \alpha_{x0}$ and $c_1 = \alpha_{y1} - \omega' \alpha_{x1}$ where $\omega = \Omega_{xx}^{-1} \omega_{xy}$; $\alpha_0 = (\alpha_{y0}, \alpha'_{x0})'$; $\alpha_1 = (\alpha_{y1}, \alpha'_{x1})'$. And $u_t = \varepsilon_{yt} - \omega_{yx} \Omega_{xx}^{-1} \varepsilon_{xt}$ with $\omega_{uu} = \omega_{yy} - \omega_{yx} \Omega_{xx}^{-1} \omega_{xy}$; note that u_t i.i.d. $N(0, \omega_{uu})$.

The matrix of long run coefficients Π has been partitioned according to:

$$\Pi = (\pi'_y, \Pi'_x)' = \alpha \beta',$$

where $\alpha = (\alpha'_{yx}, \alpha'_{xx})'$ and $\beta = (\beta'_{yx}, \beta'_{xx})'$. With $\pi_{yx.x} = \pi_{yx} - \omega' \Pi_{xx}$ (matrix $1 \times k$) and $\pi_{y.x} = \pi_y - \omega' \Pi_x$ (matrix $1 \times (k+1)$).

Matrix Π_{xx} is supposed to have rank r , $0 \leq r \leq k$ where k is the dimension of X ¹⁹. r is the minimum rank of Π and $r+1$ its maximum rank where $\Pi = \alpha \beta'$ ²⁰. When Π has rank r , one has $\pi_{yy} = 0$ as one has supposed that $\pi_{xy} = 0$ (see Pesaran *et al.* 2001) and thus:

$$\Pi = \begin{bmatrix} 0 & \pi_{yx} \\ 0 & \Pi_{xx} \end{bmatrix}.$$

In this case β_{yx} has to be null.

¹⁹ $\Pi_{xx} = \alpha_{xx} \beta'_{xx}$ where α_{xx} and β_{xx} are two matrices $k \times r$ of full column rank r .

²⁰ $\alpha = (\alpha'_{yx}, \alpha'_{xx})'$ and $\beta = (\beta'_{yx}, \beta'_{xx})'$ are two $(k+1) \times (r+1)$ dimensional matrices while $\alpha_{yx}, \alpha_{xx}, \beta_{yx}, \beta_{xx}$ are respectively $1 \times (r+1)$, $k \times (r+1)$, $1 \times (r+1)$ and $k \times (r+1)$.

B.2 Links between the parameters of the VECM and of the ones of the ARDL model

1) One notes that the long run parameters of the ARDL model are different from the ones that one would derive from a standard VECM with no contemporaneous variables ΔX_t :

$$\theta^* = -\pi_{yx}/\pi_{yy}.$$

2) It is worth emphasizing that one can not estimate the usual parameters of a VECM just from the “single equation” :

$$\Delta Y_t = c_0 + c_1 t + \pi_{yy} Y_{t-1} + \pi_{yx.x} X_{t-1} + \sum_{i=1}^{p-1} \Psi'_i \Delta Z_{t-i} + \omega' \Delta X_t + u_t.$$

Indeed, by regressing ΔY_t onto constant, time, Y_{t-1} , X_{t-1} , ΔX_t and ΔY_{t-i} , $1 \leq i \leq p-1$, one can estimate π_{yy} , $\pi_{yx.x}$, ω , and Ψ_i , $1 \leq i \leq p-1$. Accordingly, one can estimate the error-correcting intensity π_{yy} associated with a long run relationship identified by imposing that Y 's coefficient is equal to 1, but one can not estimate parameters $\pi_{yx} = \pi_{yx.x} + \omega' \Pi_{xx}$. Π_{xx} can only be estimated by jointly estimating the VECM equation of X :

$$\Delta X_t = a_0 + a_1 t + \pi_{xy} Y_{t-1} + \Pi_{xx} X_{t-1} + \sum_{i=1}^{p-1} \Gamma_{xi} \Delta Z_{t-i} + \varepsilon_{xt}.$$

C Kao's (1999) tests

C.1 The different test statistics

Let the bivariate process $w'_{it} = (\mu_{it}, \vartheta_{it})'$ for $i = 1, \dots, N$ and $t = 1, \dots, T$, with $Y_{it} = \sum_{s=1}^t \mu_{is}$

and $X_{it} = \sum_{s=1}^t \vartheta_{is}$. The long term variance-covariance Ω of w_{it} (under the homogeneity assumption) is written as:

$$\begin{aligned} \Omega &= \lim_{T \rightarrow \infty} E \left[T^{-1} \left(\sum_{t=1}^T w_{it} \right) \left(\sum_{t=1}^T w_{it} \right)' \right] \\ &= \lim_{T \rightarrow \infty} E \left[T^{-1} \sum_{t=1}^T w_{it} w'_{it} \right] + \lim_{T \rightarrow \infty} E \left[T^{-1} \sum_{k=1}^{T-1} \sum_{t=k+1}^T w_{it} w'_{it-k} \right] + \lim_{T \rightarrow \infty} E \left[T^{-1} \sum_{k=1}^{T-1} \sum_{t=k+1}^T w_{it-k} w'_{it} \right] \\ &= \Sigma + \Gamma + \Gamma' \\ \text{with } \Omega &= \begin{bmatrix} \sigma_{0\mu}^2 & \sigma_{0\mu\vartheta} \\ \sigma_{0\mu\vartheta} & \sigma_{0\vartheta}^2 \end{bmatrix}, \Sigma = \begin{bmatrix} \sigma_{\mu}^2 & \sigma_{\mu\vartheta} \\ \sigma_{\mu\vartheta} & \sigma_{\vartheta}^2 \end{bmatrix} \text{ et } \Gamma = \begin{bmatrix} \Gamma_{\mu} & \Gamma_{\mu\vartheta} \\ \Gamma_{\vartheta\mu} & \Gamma_{\vartheta} \end{bmatrix} \end{aligned}$$

From equation (A), the OLS estimator of ρ and its t-statistic are given by :

$$\hat{\rho} = \frac{\sum_{i=1}^N \sum_{t=2}^T \hat{e}_{it} \hat{e}_{it-1}}{\sum_{i=1}^N \sum_{t=2}^T \hat{e}_{it-1}^2} \quad \text{et} \quad t_{\rho} = (\hat{\rho} - 1) \sqrt{\frac{NT \sum_{i=1}^N \sum_{t=2}^T \hat{e}_{it-1}^{*2}}{\sum_{i=1}^N \sum_{t=2}^T (\hat{e}_{it}^* - \hat{\rho} \hat{e}_{it-1}^*)^2}},$$

The four Dickey-Fuller type statistics are defined as follows:

$$DF_\rho = \frac{\sqrt{NT}(\hat{\rho}-1)+3\sqrt{N}}{\sqrt{10.2}}, \quad DF_\rho^* = \frac{\sqrt{NT}(\hat{\rho}-1)+\frac{3\sqrt{N}\hat{\sigma}_v^2}{\hat{\sigma}_{0v}^2}}{\sqrt{3+\frac{36\hat{\sigma}_v^4}{5\hat{\sigma}_{0v}^4}}},$$

$$DF_t = \sqrt{1.25}t_\rho + \sqrt{1.875}N, \quad DF_t^* = \frac{t_\rho + \frac{\sqrt{6N}\hat{\sigma}_v^2}{2\hat{\sigma}_{0v}^2}}{\sqrt{\frac{\hat{\sigma}_{0v}^2}{2\hat{\sigma}_v^2} + \frac{3\hat{\sigma}_v^2}{10\hat{\sigma}_{0v}^2}}},$$

with $\sigma_{0v}^2 = \sigma_{0\mu}^2 - \sigma_{0\mu\vartheta}^2 \sigma_{0\vartheta}^{-2}$ and $\sigma_v^2 = \sigma_\mu^2 - \sigma_{\mu\vartheta}^2 \sigma_\vartheta^{-2}$ ($\hat{\sigma}_{0v}^2$ and $\hat{\sigma}_v^2$ are consistent estimators of σ_{0v}^2 and σ_v^2). The statistic of the Augmented Dickey-Fuller type test based on regression in equation (B) is:

$$ADF = \frac{t_{ADF} + \frac{\sqrt{6N}\hat{\sigma}_v}{2\hat{\sigma}_{0\mu}^2}}{\sqrt{\frac{\hat{\sigma}_{0v}^2}{2\hat{\sigma}_v^2} + \frac{3\hat{\sigma}_v^2}{10\hat{\sigma}_{0v}^2}}},$$

where t_{ADF} is the t-statistic of ρ in (B).

Kao (1999) proves according to a sequential asymptotic theory, that the DF_ρ^* , DF_t^* and ADF_t statistics follow a $N(0, 1)$ distribution. However, results from various Monte Carlo simulations by cannot conclude to the superiority of one statistic, since the results are very sensitive to the Data Generating Process. In our study, we refer to econometric theory and focus on the statistics whose distribution is not affected by nuisance parameters, namely DF_ρ^* , DF_t^* and ADF_t .

C.2 The augmented regression in the ADF test

The cointegration test proposed by Kao is a unit root test for the residuals \hat{e}_{it} of the long run equation:

$$\Delta \hat{e}_{it} = \rho \hat{e}_{it-1} + v_{it}$$

which can be rewritten as:

$$\begin{aligned} \Delta \left(\tilde{Y}_{it} - \tilde{X}'_{it} \hat{\beta} \right) &= \rho \left(\tilde{Y}_{it-1} - \tilde{X}'_{it-1} \hat{\beta} \right) + v_{it} \\ \Leftrightarrow \Delta \left(\tilde{Y}_{it} \right) &= \rho \left(\tilde{Y}_{it-1} - \tilde{X}'_{it} \hat{\beta} \right) + (\rho + 1) \Delta \tilde{X}'_{it} \hat{\beta} + v_{it} \end{aligned}$$

To get the ADF statistic, the previous regression is augmented as:

$$\begin{aligned} \hat{e}_{it} &= \rho \hat{e}_{it-1} + \sum_{j=1}^p \varphi_j \Delta \hat{e}_{it-j} + v_{itp}, \\ \Leftrightarrow \Delta \left(\tilde{Y}_{it} \right) &= \rho \left(\tilde{Y}_{it-1} - \tilde{X}'_{it-1} \hat{\beta} \right) + \Delta \tilde{X}'_{it} \hat{\beta} + \sum_{j=1}^p \varphi_j \Delta \left(\tilde{Y}_{it-j} - \tilde{X}'_{it-j} \hat{\beta} \right) + v_{itp} \\ \Leftrightarrow \Delta \left(\tilde{Y}_{it} \right) &= \rho \left(\tilde{Y}_{it-1} - \tilde{X}'_{it} \hat{\beta} \right) + (\rho + 1) \Delta \tilde{X}'_{it} \hat{\beta} + \sum_{j=1}^p \varphi_j \Delta \tilde{Y}_{it-j} + \sum_{j=1}^p (-\varphi_j) \Delta \tilde{X}'_{it-j} \hat{\beta} + v_{itp} \end{aligned}$$

which appears to be a constrained version of the regression implemented in the ARDL framework (see Reqr. 2).

Notes d'Études et de Recherche

1. C. Huang and H. Pagès, "Optimal Consumption and Portfolio Policies with an Infinite Horizon: Existence and Convergence," May 1990.
2. C. Bordes, « Variabilité de la vitesse et volatilité de la croissance monétaire : le cas français », février 1989.
3. C. Bordes, M. Driscoll and A. Sauviat, "Interpreting the Money-Output Correlation: Money-Real or Real-Real?," May 1989.
4. C. Bordes, D. Goyeau et A. Sauviat, « Taux d'intérêt, marge et rentabilité bancaires : le cas des pays de l'OCDE », mai 1989.
5. B. Bensaïd, S. Federbusch et R. Gary-Bobo, « Sur quelques propriétés stratégiques de l'intéressement des salariés dans l'industrie », juin 1989.
6. O. De Bandt, « L'identification des chocs monétaires et financiers en France : une étude empirique », juin 1990.
7. M. Boutillier et S. Dérangère, « Le taux de crédit accordé aux entreprises françaises : coûts opératoires des banques et prime de risque de défaut », juin 1990.
8. M. Boutillier and B. Cabrillac, "Foreign Exchange Markets: Efficiency and Hierarchy," October 1990.
9. O. De Bandt et P. Jacquinet, « Les choix de financement des entreprises en France : une modélisation économétrique », octobre 1990 (English version also available on request).
10. B. Bensaïd and R. Gary-Bobo, "On Renegotiation of Profit-Sharing Contracts in Industry," July 1989 (English version of NER n° 5).
11. P. G. Garella and Y. Richelle, "Cartel Formation and the Selection of Firms," December 1990.
12. H. Pagès and H. He, "Consumption and Portfolio Decisions with Labor Income and Borrowing Constraints," August 1990.
13. P. Sicsic, « Le franc Poincaré a-t-il été délibérément sous-évalué ? », octobre 1991.
14. B. Bensaïd and R. Gary-Bobo, "On the Commitment Value of Contracts under Renegotiation Constraints," January 1990 revised November 1990.
15. B. Bensaïd, J.-P. Lesne, H. Pagès and J. Scheinkman, "Derivative Asset Pricing with Transaction Costs," May 1991 revised November 1991.
16. C. Monticelli and M.-O. Strauss-Kahn, "European Integration and the Demand for Broad Money," December 1991.
17. J. Henry and M. Phelipot, "The High and Low-Risk Asset Demand of French Households: A Multivariate Analysis," November 1991 revised June 1992.
18. B. Bensaïd and P. Garella, "Financing Takeovers under Asymmetric Information," September 1992.

19. A. de Palma and M. Uctum, "Financial Intermediation under Financial Integration and Deregulation," September 1992.
20. A. de Palma, L. Leruth and P. Régibeau, "Partial Compatibility with Network Externalities and Double Purchase," August 1992.
21. A. Frachot, D. Janci and V. Lacoste, "Factor Analysis of the Term Structure: a Probabilistic Approach," November 1992.
22. P. Sicsic et B. Villeneuve, « L'afflux d'or en France de 1928 à 1934 », janvier 1993.
23. M. Jeanblanc-Picqué and R. Avesani, "Impulse Control Method and Exchange Rate," September 1993.
24. A. Frachot and J.-P. Lesne, "Expectations Hypothesis and Stochastic Volatilities," July 1993 revised September 1993.
25. B. Bensaid and A. de Palma, "Spatial Multiproduct Oligopoly," February 1993 revised October 1994.
26. A. de Palma and R. Gary-Bobo, "Credit Contraction in a Model of the Banking Industry," October 1994.
27. P. Jacquinet et F. Mihoubi, « Dynamique et hétérogénéité de l'emploi en déséquilibre », septembre 1995.
28. G. Salmat, « Le retournement conjoncturel de 1992 et 1993 en France : une modélisation VAR », octobre 1994.
29. J. Henry and J. Weidmann, "Asymmetry in the EMS Revisited: Evidence from the Causality Analysis of Daily Eurorates," February 1994 revised October 1994.
30. O. De Bandt, "Competition Among Financial Intermediaries and the Risk of Contagious Failures," September 1994 revised January 1995.
31. B. Bensaid et A. de Palma, « Politique monétaire et concurrence bancaire », janvier 1994 révisé en septembre 1995.
32. F. Rosenwald, « Coût du crédit et montant des prêts : une interprétation en terme de canal large du crédit », septembre 1995.
33. G. Cette et S. Mahfouz, « Le partage primaire du revenu : constat descriptif sur longue période », décembre 1995.
34. H. Pagès, "Is there a Premium for Currencies Correlated with Volatility? Some Evidence from Risk Reversals," January 1996.
35. E. Jondeau and R. Ricart, "The Expectations Theory: Tests on French, German and American Euro-rates," June 1996.
36. B. Bensaid et O. De Bandt, « Les stratégies "stop-loss" : théorie et application au Contrat Notionnel du Matif », juin 1996.
37. C. Martin et F. Rosenwald, « Le marché des certificats de dépôts. Écarts de taux à l'émission : l'influence de la relation émetteurs-souscripteurs initiaux », avril 1996.

38. Banque de France - CEPREMAP - Direction de la Prévision - Erasme - INSEE - OFCE, « Structures et propriétés de cinq modèles macroéconomiques français », juin 1996.
39. F. Rosenwald, « L'influence des montants émis sur le taux des certificats de dépôts », octobre 1996.
40. L. Baumel, « Les crédits mis en place par les banques AFB de 1978 à 1992 : une évaluation des montants et des durées initiales », novembre 1996.
41. G. Cette et E. Kremp, « Le passage à une assiette valeur ajoutée pour les cotisations sociales : Une caractérisation des entreprises non financières “gagnantes” et “perdantes” », novembre 1996.
42. S. Avouyi-Dovi, E. Jondeau et C. Lai Tong, « Effets “volume”, volatilité et transmissions internationales sur les marchés boursiers dans le G5 », avril 1997.
43. E. Jondeau et R. Ricart, « Le contenu en information de la pente des taux : Application au cas des titres publics français », juin 1997.
44. B. Bensaid et M. Boutillier, « Le contrat notionnel : efficience et efficacité », juillet 1997.
45. E. Jondeau et R. Ricart, « La théorie des anticipations de la structure par terme : test à partir des titres publics français », septembre 1997.
46. E. Jondeau, « Représentation VAR et test de la théorie des anticipations de la structure par terme », septembre 1997.
47. E. Jondeau et M. Rockinger, « Estimation et interprétation des densités neutres au risque : Une comparaison de méthodes », octobre 1997.
48. L. Baumel et P. Sevestre, « La relation entre le taux de crédits et le coût des ressources bancaires. Modélisation et estimation sur données individuelles de banques », octobre 1997.
49. P. Sevestre, “On the Use of Banks Balance Sheet Data in Loan Market Studies : A Note,” October 1997.
50. P.-C. Hautcoeur and P. Sicsic, “Threat of a Capital Levy, Expected Devaluation and Interest Rates in France during the Interwar Period,” January 1998.
51. P. Jacquinot, « L'inflation sous-jacente à partir d'une approche structurelle des VAR : une application à la France, à l'Allemagne et au Royaume-Uni », janvier 1998.
52. C. Bruneau et O. De Bandt, « La modélisation VAR structurel : application à la politique monétaire en France », janvier 1998.
53. C. Bruneau and E. Jondeau, “Long-Run Causality, with an Application to International Links between Long-Term Interest Rates,” June 1998.
54. S. Coutant, E. Jondeau and M. Rockinger, “Reading Interest Rate and Bond Futures Options’ Smiles: How PIBOR and Notional Operators Appreciated the 1997 French Snap Election,” June 1998.
55. E. Jondeau et F. Sédillot, « La prévision des taux longs français et allemands à partir d'un modèle à anticipations rationnelles », juin 1998.

56. E. Jondeau and M. Rockinger, "Estimating Gram-Charlier Expansions with Positivity Constraints," January 1999.
57. S. Avouyi-Dovi and E. Jondeau, "Interest Rate Transmission and Volatility Transmission along the Yield Curve," January 1999.
58. S. Avouyi-Dovi et E. Jondeau, « La modélisation de la volatilité des bourses asiatiques », janvier 1999.
59. E. Jondeau, « La mesure du ratio rendement-risque à partir du marché des euro-devises », janvier 1999.
60. C. Bruneau and O. De Bandt, "Fiscal Policy in the Transition to Monetary Union: A Structural VAR Model," January 1999.
61. E. Jondeau and R. Ricart, "The Information Content of the French and German Government Bond Yield Curves: Why Such Differences?," February 1999.
62. J.-B. Chatelain et P. Sevestre, « Coûts et bénéfices du passage d'une faible inflation à la stabilité des prix », février 1999.
63. D. Irac et P. Jacquinot, « L'investissement en France depuis le début des années 1980 », avril 1999.
64. F. Mihoubi, « Le partage de la valeur ajoutée en France et en Allemagne », mars 1999.
65. S. Avouyi-Dovi and E. Jondeau, "Modelling the French Swap Spread," April 1999.
66. E. Jondeau and M. Rockinger, "The Tail Behavior of Stock Returns: Emerging Versus Mature Markets," June 1999.
67. F. Sédillot, « La pente des taux contient-elle de l'information sur l'activité économique future ? », juin 1999.
68. E. Jondeau, H. Le Bihan et F. Sédillot, « Modélisation et prévision des indices de prix sectoriels », septembre 1999.
69. H. Le Bihan and F. Sédillot, "Implementing and Interpreting Indicators of Core Inflation: The French Case," September 1999.
70. R. Lacroix, "Testing for Zeros in the Spectrum of an Univariate Stationary Process: Part I," December 1999.
71. R. Lacroix, "Testing for Zeros in the Spectrum of an Univariate Stationary Process: Part II," December 1999.
72. R. Lacroix, "Testing the Null Hypothesis of Stationarity in Fractionally Integrated Models," December 1999.
73. F. Chesnay and E. Jondeau, "Does correlation between stock returns really increase during turbulent period?," April 2000.
74. O. Burkart and V. Coudert, "Leading Indicators of Currency Crises in Emerging Economies," May 2000.
75. D. Irac, "Estimation of a Time Varying NAIRU for France," July 2000.

76. E. Jondeau and H. Le Bihan, "Evaluating Monetary Policy Rules in Estimated Forward-Looking Models: A Comparison of US and German Monetary Policies," October 2000.
77. E. Jondeau and M. Rockinger, "Conditional Volatility, Skewness, and Kurtosis: Existence and Persistence," November 2000.
78. P. Jacquinot et F. Mihoubi, « Modèle à Anticipations Rationnelles de la CONjoncture Simulée : MARCOS », novembre 2000.
79. M. Rockinger and E. Jondeau, "Entropy Densities: With an Application to Autoregressive Conditional Skewness and Kurtosis," January 2001.
80. B. Amable and J.-B. Chatelain, "Can Financial Infrastructures Foster Economic Development?," January 2001.
81. J.-B. Chatelain and J.-C. Teurlai, "Pitfalls in Investment Euler Equations," January 2001.
82. M. Rockinger and E. Jondeau, "Conditional Dependency of Financial Series: An Application of Copulas," February 2001.
83. C. Florens, E. Jondeau and H. Le Bihan, "Assessing GMM Estimates of the Federal Reserve Reaction Function," March 2001.
84. J.-B. Chatelain, "Mark-up and Capital Structure of the Firm facing Uncertainty," June 2001.
85. B. Amable, J.-B. Chatelain and O. De Bandt, "Optimal Capacity in the Banking Sector and Economic Growth," June 2001.
86. E. Jondeau and H. Le Bihan, "Testing for a Forward-Looking Phillips Curve. Additional Evidence from European and US Data," December 2001.
87. G. Clette, J. Mairesse et Y. Kocoglu, « Croissance économique et diffusion des TIC : le cas de la France sur longue période (1980-2000) », décembre 2001.
88. D. Irac and F. Sédillot, "Short Run Assessment of French Economic Activity Using OPTIM," January 2002.
89. M. Baghli, C. Bouthevillain, O. de Bandt, H. Fraisse, H. Le Bihan et Ph. Rousseaux, « PIB potentiel et écart de PIB : quelques évaluations pour la France », juillet 2002.
90. E. Jondeau and M. Rockinger, "Asset Allocation in Transition Economies," October 2002.
91. H. Pagès and J.A.C. Santos, "Optimal Supervisory Policies and Depositor-Preferences Laws," October 2002.
92. C. Loupias, F. Savignac and P. Sevestre, "Is There a Bank Lending Channel in France? Evidence from Bank Panel Data," November 2002.
93. M. Ehrmann, L. Gambacorta, J. Martínez-Pagés, P. Sevestre and A. Worms, "Financial Systems and The Role in Monetary Policy Transmission in the Euro Area," November 2002.
94. S. Avouyi-Dovi, D. Guégan et S. Ladoucette, « Une mesure de la persistance dans les indices boursiers », décembre 2002.

95. S. Avouyi-Dovi, D. Guégan et S. Ladoucette, "What is the Best Approach to Measure the Interdependence between Different Markets? ," December 2002.
96. J.-B. Chatelain and A. Tiomo, "Investment, the Cost of Capital and Monetary Policy in the Nineties in France: A Panel Data Investigation," December 2002.
97. J.-B. Chatelain, A. Generale, I. Hernando, U. von Kalckreuth and P. Vermeulen, "Firm Investment and Monetary Policy Transmission in the Euro Area," December 2002.
98. J.-S. Mésonnier, « Banque centrale, taux de l'escompte et politique monétaire chez Henry Thornton (1760-1815) », décembre 2002.
99. M. Baghli, G. Cette et A. Sylvain, « Les déterminants du taux de marge en France et quelques autres grands pays industrialisés : Analyse empirique sur la période 1970-2000 », janvier 2003.
100. G. Cette and Ch. Pfister, "The Challenges of the "New Economy" for Monetary Policy," January 2003.
101. C. Bruneau, O. De Bandt, A. Flageollet and E. Michaux, "Forecasting Inflation using Economic Indicators: the Case of France," May 2003.
102. C. Bruneau, O. De Bandt and A. Flageollet, "Forecasting Inflation in the Euro Area," May 2003.
103. E. Jondeau and H. Le Bihan, "ML vs GMM Estimates of Hybrid Macroeconomic Models (With an Application to the "New Phillips Curve")," September 2003.
104. J. Matheron and T.-P. Maury, "Evaluating the Fit of Sticky Price Models," January 2004.
105. S. Moyen and J.-G. Sahuc, "Incorporating Labour Market Frictions into an Optimising-Based Monetary Policy Model," January 2004.
106. M. Baghli, V. Brunhes-Lesage, O. De Bandt, H. Fraise et J.-P. Villetelle, « MASCOTTE : Modèle d'Analyse et de préviSion de la COnjoncture TrimesTrielle », février 2004.
107. E. Jondeau and M. Rockinger, "The Bank Bias: Segmentation of French Fund Families," February 2004.
108. E. Jondeau and M. Rockinger, "Optimal Portfolio Allocation Under Higher Moments," February 2004.
109. C. Bordes et L. Clerc, « Stabilité des prix et stratégie de politique monétaire unique », mars 2004.
110. N. Belorgey, R. Lecat et T.-P. Maury, « Déterminants de la productivité par employé : une évaluation empirique en données de panel », avril 2004.
111. T.-P. Maury and B. Pluyaud, "The Breaks in per Capita Productivity Trends in a Number of Industrial Countries," April 2004.
112. G. Cette, J. Mairesse and Y. Kocoglu, "ICT Diffusion and Potential Output Growth," April 2004.
113. L. Baudry, H. Le Bihan, P. Sevestre and S. Tarrieu, "Price Rigidity. Evidence from the French CPI Micro-Data," September 2004.

114. C. Bruneau, O. De Bandt and A. Flageollet, "Inflation and the Markup in the Euro Area," September 2004.
115. J.-S. Mésonnier and J.-P. Renne, "A Time-Varying "Natural" Rate of Interest for the Euro Area," September 2004.
116. G. Cette, J. Lopez and P.-S. Noual, "Investment in Information and Communication Technologies: an Empirical Analysis," October 2004.
117. J.-S. Mésonnier et J.-P. Renne, « Règle de Taylor et politique monétaire dans la zone euro », octobre 2004.
118. J.-G. Sahuc, "Partial Indexation, Trend Inflation, and the Hybrid Phillips Curve," December 2004.
119. C. Loupias et B. Wigniolle, « Régime de retraite et chute de la natalité : évolution des mœurs ou arbitrage micro-économique ? », décembre 2004.
120. C. Loupias and R. Ricart, "Price Setting in France: new Evidence from Survey Data," December 2004.
121. S. Avouyi-Dovi and J. Matheron, "Interactions between Business Cycles, Stock Markets Cycles and Interest Rates: the Stylised Facts," January 2005.
122. L. Bilke, "Break in the Mean and Persistence of Inflation: a Sectoral Analysis of French CPI," January 2005.
123. S. Avouyi-Dovi and J. Matheron, "Technology Shocks and Monetary Policy in an Estimated Sticky Price Model of the US Economy," April 2005.
124. M. Dupaigne, P. Fève and J. Matheron, "Technology Shock and Employment: Do We Really Need DSGE Models with a Fall in Hours?," June 2005.
125. P. Fève and J. Matheron, "Can the Kydland-Prescott Model Pass the Cogley-Nason Test?," June 2005.
126. S. Avouyi-Dovi and J. Matheron, "Technology Shocks and Monetary Policy in an Estimated Sticky Price Model of the Euro Area," June 2005.
127. O. Loisel, "Central Bank Reputation in a Forward-Looking Model," June 2005.
128. B. Bellone, E. Gautier et S. Le Coent, « Les marchés financiers anticipent-ils les retournements conjoncturels ? », juillet 2005.
129. P. Fève, « La modélisation macro-économétrique dynamique », juillet 2005.
130. G. Cette, N. Dromel and D. Méda, "Opportunity Costs of Having a Child, Financial Constraints and Fertility," August 2005.
131. S. Gouteron et D. Szpiro, « Excès de liquidité monétaire et prix des actifs », septembre 2005.
132. J. Baude, « L'impact des chocs boursiers sur le crédit en France depuis le milieu des années quatre-vingt-dix », septembre 2005.
133. R. Boulès and G. Cette, "A Comparison of Structural Productivity Levels in the Major Industrialised Countries," October 2005.

134. T. Grunspan, "The Fed and the Question of Financial Stability: An Empirical Investigation," October 2005.
135. S. Fabiani, M. Druant, I. Hernando, C. Kwapil, B. Landau, C. Loupias, F. Martins, T. Mathä, R. Sabbatini, H. Stahl and A. Stockman, "The Pricing Behaviour of Firms in the Euro Area: New Survey Evidence," November 2005.
136. E. Dhyne, L. Alvarez, H. Le Bihan, G. Veronese, D. Dias, J. Hoffmann, N. Jonker, P. Lünemann, F. Ruml and J. Vilminen, "Price Setting in the Euro Area: Some Stylized Facts from Individual Consumer Price Data," November 2005.
137. D. Fougère, H. Le Bihan and P. Sevestre, "Heterogeneity in Consumer Price Stickiness: A Microeconomic Investigation," November 2005.
138. L. Alvarez, E. Dhyne, M. Hoeberichts, C. Kwapil, H. Le Bihan, P. Lünemann, F. Martins, R. Sabbatini, H. Stahl, P. Vermeulen and J. Vilminen, "Sticky Prices in the Euro Area: a Summary of New Micro Evidence," November 2005.
139. E. Kharroubi, "Illiquidity, Financial Development and the Growth-Volatility Relationship," February 2006.
140. M. Baghli, C. Cahn and H. Fraise, "Is the Inflation-Output Nexus Asymmetric in the Euro Area," April 2006.
141. E. Jondeau and J-G. Sahuc, "Optimal Monetary Policy in an Estimated DSGE Model of the Euro Area with Cross-country Heterogeneity," April 2006.
142. S. Avouyi-Dovi, M. Brun, A. Dreyfus, F. Drumetz, V. Oung et J.-G. Sahuc, « La fonction de demande de monnaie pour la zone euro : un réexamen », mai 2006.
143. C. Jardet, "Term Structure Anomalies : Term Premium or Peso Problem?" May 2006.
144. S. Avouyi-Dovi, R. Kierzenkowski and C. Lubochinsky, "Are Business and Credit Cycles Converging or Diverging? A comparison of Poland, Hungary, the Czech Republic and the Euro Area", May 2006.
145. O. De Bandt, C. Bruneau and A. Flageollet, "Assessing Aggregate Comovements in France, Germany and Italy. Using a Non Stationary Factor Model of the Euro Area" June 2006.
146. M. Baghli, C. Cahn and J-P. Villette, "Estimating Potential Output with a Production Function for France, Germany and Italy", June 2006.
147. E. Fonteny, « La désaisonnalisation des séries d'agrégats monétaires et de crédit à la Banque de France : aspects théoriques et mise en œuvre », juin 2006.
148. J. Matheron and C. Poilly, "How Well Does a Small Structural Model with Sticky Prices and Wages Fit Postwar U.S. Data", July 2006.
149. E. Kharroubi, "Financial (Dis)Integration," July 2006.
150. J. Carrillo, P. Fève and J. Matheron, "Monetary Policy Inertia or Persistent Shocks?" July 2006.
151. C. Ewerhart, N. Cassola and N. Valla, "Declining Valuations and Equilibrium Bidding in Central Bank Refinancing Operations," August 2006.

152. D. Fougère, « Réformes structurelles sur le marché du travail : quels enseignements peut-on tirer des études existantes », août 2006.
153. D. Irac, “Revisiting the proximity-concentration trade-off: Distance and Horizontal Foreign Direct Investment in OECD Countries,” August 2006.
154. D. Irac and C. Minoiu, “Risk insurance in a Transition Economy: Evidence from Rural Romania,” August 2006.
155. H. Lustig and A. Verdelhan, “The Cross-Section of Foreign Currency Risk Premia and Consumption Growth Risk,” August 2006.
156. R. Boursès and G. Cette, “Trends in “structural” productivity levels in the major industrialized countries,” September 2006.
157. J.-S. Mésonnier, “The Reliability of Macroeconomic Forecasts based on Real Interest Rate Gap Estimates in Real Time: an Assessment for the Euro Area,” October 2006
158. O. de Bandt, C. Bruneau and W. El Amri, “Convergence in Household Credit Demand Across Euro Area Countries: Evidence from Panel Data,” October 2006

Pour tous commentaires ou demandes sur les Notes d'Études et de Recherche, contacter la bibliothèque de la direction de la recherche à l'adresse suivante :

For any comment or enquiries on the Working Papers, contact the library of the Research Directorate at the following address :

BANQUE DE FRANCE
41- 1404 Labolog
75049 Paris Cedex 01
tél : 0033 (0)1 42 92 49 55 ou 62 65
fax :0033 (0)1 42 92 62 92
email : thierry.demoulin@banque-france.fr
jeannine.agoutin@banque-france.fr