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Deep Roots of Fiscal Behavior

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ABSTRACT

This paper investigates the determinants of fiscal policy behavior and its time-varying volatility, using panel data for a broad set of advanced and emerging market economies during the period 1990–2012. The empirical results show that discretionary fiscal policy is influenced by policy inertia, the level of public debt, and the output gap in both advanced and emerging-market economies. In addition, the paper finds that macro-financial factors (such as real exchange rate, financial development, interest rates, asset prices, and natural resource rents) and demographic and institutional factors (such as the old-age dependency ratio, the quality of institutions, and policy anchors such as fiscal rules and IMF-supported stabilization programs) tend to have a significant effect on fiscal policy behavior. The results also indicate that higher government debt leads to more volatile fiscal behavior, while fiscal rules and higher institutional quality reduce the volatility of fiscal policy over time.

JEL classification: E60, E62, G01, H30, H62

Keywords: Fiscal policy, fiscal reaction functions, fiscal policy volatility

1. INTRODUCTION

The 2008–09 global economic crisis has led to an unprecedented erosion of fiscal positions in both advanced and emerging-market economies across the world. The average fiscal deficit in advanced economies rose from 1.1 percent of GDP in 2007 to 9 percent in 2009, while the increase was smaller but still sizable in emerging-market economies, from a surplus of 1.3 percent to a deficit of 4.6 percent. Although budget deficits narrowed by 2012, fiscal vulnerabilities remain elevated, especially in advanced economies with gross public debt levels of 110 percent of GDP,

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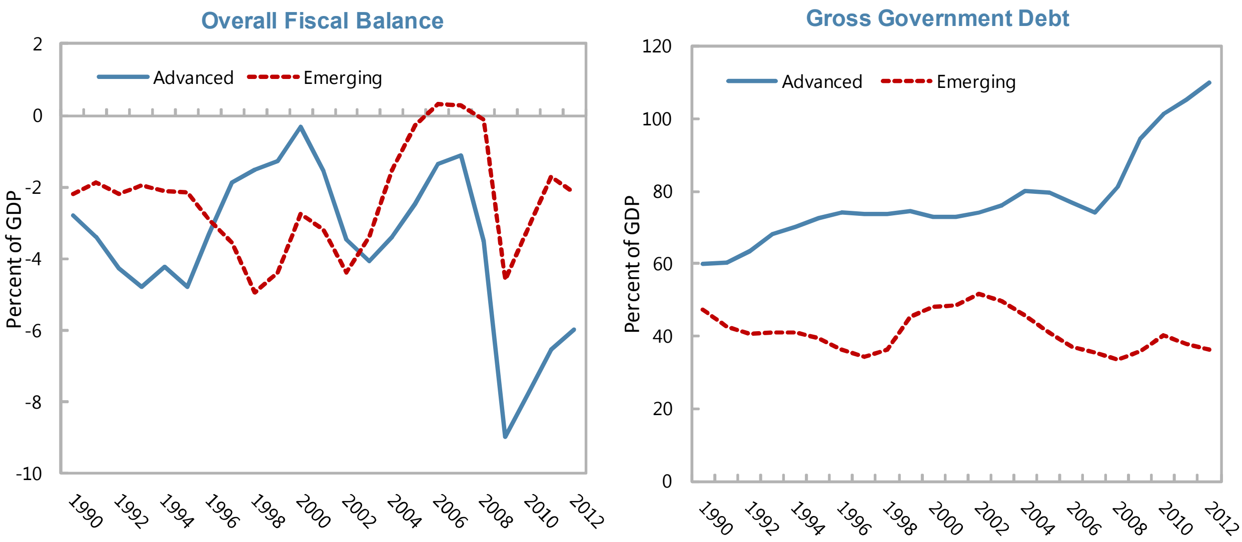
up, on average, from 74 percent in 2007. Furthermore, fiscal policy has become more volatile, with the standard deviation of the overall budget balance rising from an average of 0.9 between 2000 and 2007 to 1.7 after the crisis in advanced economies and from 1.0 to 1.2 in emerging markets, partly reflecting more the volatile environment. Although these unusual developments have generally been blamed on fiscal stimulus packages, the severity of the economic downturn, and the bailout of financial institutions, the extent of the change in fiscal positions and the variation in policy responses call for an in-depth analysis of the underlying determinants of discretionary fiscal policy and its volatility over time.

Figure 1
Budget Balance and Government Debt, 1990–2012



Source: International Monetary Fund; Authors' calculations.

An extensive literature on fiscal reaction functions links fiscal behavior to a host of macro-financial and institutional characteristics. Most empirical studies find that the fiscal-policy stance tends to be procyclical, although theoretical considerations require neutral or countercyclical behavior. The literature has identified a link between cross-country divergence in discretionary fiscal policy and a wide range of macro-financial and demographic characteristics and institutional and political features. There is also empirical evidence suggesting that fiscal rules and external anchors, such as an International Monetary Fund (IMF)-supported economic stabilization programs, tend to influence fiscal behavior. Even though empirical findings are broadly consistent, most of the literature focuses on advanced economies and covers the period prior to the 2008–09 economic crisis. These studies also tend to rely on unbalanced panels and do not deal with econometric complications associated with endogenous regressors and country-specific heterogeneity.

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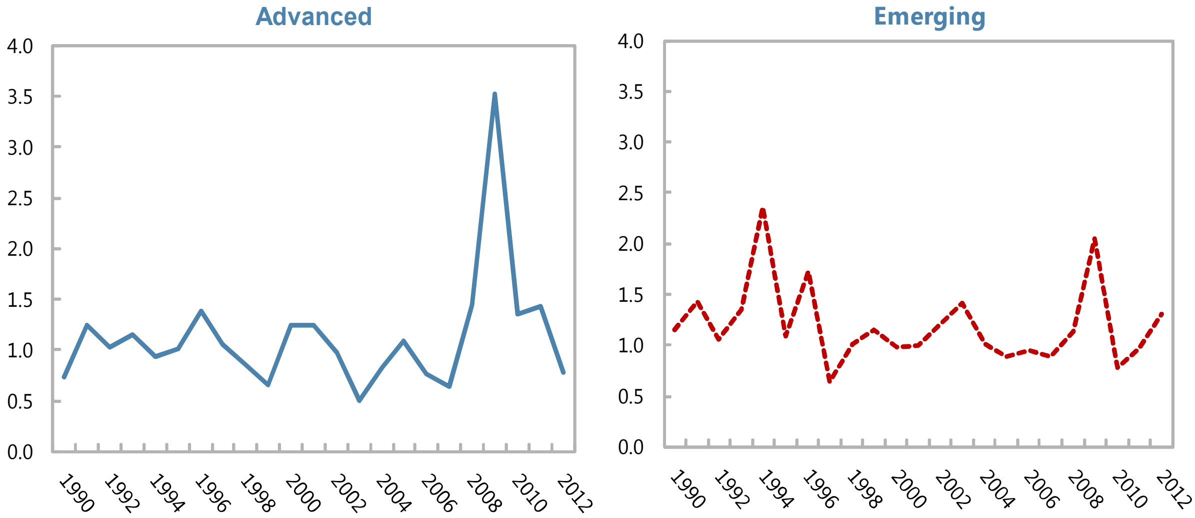
Figure 2
Fiscal Policy Volatility, 1990–2012



Source: International Monetary Fund; Authors' calculations.

This paper empirically investigates the main determinants of discretionary fiscal policy and its volatility in advanced and emerging-market economies. Firstly, and differently from previous studies, this paper estimates fiscal reaction functions for a large panel of advanced and emerging-market economies, using the cyclically-adjusted primary budget balance as a measure of discretionary fiscal policy and focusing on possible differences between the two groups of countries.[2] Secondly, this paper studies a wide range of economic, financial, demographic and institutional variables as potential determinants over two decades including the aftermath of the 2008–09 crisis. Thirdly, as the crisis has shown, macroeconomic volatility can have larger impact than previously expected, and therefore this paper estimates the determinants of time-varying volatility of discretionary fiscal policy, as measured by the standard deviation of the cyclically-adjusted primary balance. Fourthly, we adopt a dynamic panel estimation approach based on the Generalized Method of Moments (GMM) that corrects for biases associated with endogenous regressors and unobserved country-specific heterogeneity. Finally, this paper relies on a balanced panel that yields more robust estimations than the unbalanced panels widely used in the literature.[3]

The empirical results suggest that discretionary fiscal policy is influenced by a range of macro-financial and institutional factors. In line with the literature, the results show that discretionary fiscal policy exhibits a high degree of inertia and responds to the level of public debt and the output gap in both advanced and emerging-market economies. In addition, we find that macro-financial factors (such as the real exchange rate, financial development, interest rates, asset prices, and natural resource rents) and demographic and institutional factors (such as the old-age dependency ratio, the quality of institutions, and policy anchors such as fiscal rules and IMF-supported stabilization programs) tend to have a significant influence fiscal behavior. The results indicate that these factors tend to influence discretionary fiscal policy differently in advanced and emerging-market economies and that policy inertia is significantly greater in the post-crisis period. We also find strong evidence that fiscal policy behavior after the recent crisis has turned even more procyclical and become less responsive to the government's intertemporal budget constraint and, therefore, long-run fiscal solvency concerns. Finally, we show that higher government debt

[2] Although there are some caveats in estimation, the cyclically-adjusted primary balance helps minimize the simultaneity bias that may arise as policy decisions and growth interact. For a criticism of the cyclically-adjusted primary budget balance as an approximation of discretionary fiscal policy, see Mélitz (2000), Alberola, Mínguez, De Cos, and Marqués (2003), Larch and Salto (2005), and Riera-Crichton, Végh, and Vuletin (2012).

[3] Balanced panels provide equal heterogeneity conditional distribution, avoid the initial value problem with observations entering at the same time points, and produce consistent estimates.

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leads to more volatile policy behavior, while fiscal rules and higher institutional quality reduce fiscal volatility.

The remainder of this paper is organized as follows. Section II provides a brief overview of the literature, followed by a summary of panel data sources in Section III. Sections IV and V describe our empirical strategy and the estimation results, respectively. Robustness checks are presented in Section VI, while Sections VII offers some concluding remarks.

2. THEORATICAL AND EMPIRICAL BACKGROUND

The tax-smoothing model with perfect foresight proposed by Barro (1979) and Lucas and Stokey (1983) argues that fiscal policy is determined by the government's need to smooth distortions associated with taxation. Accordingly, revenue and spending shocks should be absorbed by budget deficits during economic recessions and by surpluses in times of economic expansion. From an empirical point of view, however, the tax-smoothing theory cannot explain the persistence of budget deficits, and why countries facing similar economic shocks experience in reality different fiscal policy paths. Many empirical studies find that the fiscal policy stance tends to be procyclical, contrary to theoretical considerations calling for neutral or countercyclical behavior.[4] However, most of the existing literature does not use a cyclically-adjusted measure of the fiscal policy stance and tends to rely on unbalanced panels with a focus on some variables of particular interest.

The empirical literature has identified a link between cross-country divergence in fiscal behavior and a range of macro-financial and demographic characteristics. Easterly and Rebelo (1993) found evidence that the budget balance is mainly correlated with economic growth, as well as with the level of per capita income, leading to diverging fiscal trends between countries at different stages of development. Several studies have focused on the impact of interest rates in modeling fiscal policy behavior (Baldacci et al., 2011 and Kiley, 2012), while others have identified trade openness, financial development and natural resource rents as important factors in determining fiscal policy behavior (Alesina and Perotti, 1995). The fiscal consequences of financial crises are also considered to have a significant effect on discretionary policy decisions (Schaechter et al., 2012). Demographic characteristics also appear to influence fiscal policy behavior (Woo, 2003, 2009).

Indeed, political institutions appear to play a role in determining the extent and persistence of fiscal imbalances. Building on the theory of political business cycles, Roubini and Sachs (1989) showed that government fragmentation tends to result in large and persistent budget deficits and excessive fiscal policy reactions in response to economic shocks. Other empirical studies have confirmed this relationship between fiscal performance and a wide range of institutional and political factors, including budget institutions and procedures, type of political regime, government structure, ideological orientation, electoral cycles, the quality of institutions, and corruption, among other things, in samples of advanced and developing countries.[5] Some papers also find that fiscal rules and external anchors, such as an IMF-supported program, tend to have a positive effect on fiscal policy behavior (see, for example, Celasun, Debrun, and Ostry, 2006; Debrun et al., 2008; Poplawski-Ribeiro, 2009; Ghosh et al., 2013).

An important aspect of fiscal policy behavior is its time-varying volatility, which may have significant macro-financial implications. Although it has received less attention in the literature,

[4] See, for example, Gavin and Perotti, 1997; Bohn, 1998; Talvi and Végh, 2000; Favero, 2002; Galí and Perotti, 2003; Lane, 2003; Balassone and Francese, 2004; Kaminsky, Reinhart, and Végh, 2004; Alesina and Tabellini, 2005; Annett, 2006; Wyplosz, 2006; Debrun and Kumar, 2007; Ilzetzki, Mendoza and Végh, 2010.

[5] Eslava (2011) provides a recent overview of the literature on the political economy of fiscal policy.

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excessive volatility in fiscal policy can undermine fiscal sustainability and lead to macro-financial distortions. Ramey and Ramey (1995) presented evidence that government spending volatility has a negative effect on real GDP per capita growth. Looking at the underlying determinants, Furceri and Poplawski-Riberio (2008) found that smaller countries tend to have more volatile government spending, while Agnello and Sousa (2009) observed significant linkages between deficit volatility and the level of economic development, political instability, and inflation, especially in countries with more trade openness. From a macro-fiscal point of view, Fatás and Mihov (2003) showed that numerical fiscal rules tend to lead to a lower degree of volatility in fiscal policy implementation.

This paper contributes to the empirical literature on the determinants of fiscal reaction functions in four ways. Firstly, and differently from most of the existing literature, we use the cyclically-adjusted primary budget balance as a measure of discretionary fiscal policy in a balanced panel of advanced and emerging market economies over two decades including the aftermath of the 2008–09 crisis. Secondly, we include a comprehensive set of economic, financial, demographic and institutional variables as potential determinants of fiscal policy behavior. Thirdly, we augment the analysis of fiscal behavior by estimating the determinants of time-varying volatility of discretionary fiscal policy, as proxied by the standard deviation of the cyclically-adjusted primary balance. Fourthly, we adopt a dynamic panel estimation approach that corrects for biases associated with endogenous regressors and unobserved country-specific heterogeneity.

3. DATA

We construct a panel dataset covering the period 1990–2012 and consisting of 49 advanced and emerging-market economies.[6] The panel includes 24 advanced and 25 emerging market economies (Table A1). While the focus is on two groups of countries, we do not explore differences across regions or the influence of regional factors on fiscal policy. The dependent variable – the cyclically-adjusted primary budget balance – is based on the IMF's Public Finances in Modern History database, assembled from historical sources by Mauro et al. (2013).[7] We measure the volatility of fiscal policy behavior as the standard deviation of the cyclically-adjusted primary budget balance over two years, and estimate the output gap for each country by applying the Hodrick-Prescott (HP) filter to decompose real GDP into trend and cyclical components (Hodrick and Prescott, 1997).

Economic and financial series are drawn from the IMF's International Financial Statistics (IFS) and World Economic Outlook (WEO) databases and the World Bank's World Development Indicators (WDI) database. Data on institutional and political indicators come from the World Bank's Database of Political Institutions (DPI), the Polity IV Database, and the International Country Risk Guide. The crisis dummy – taking the value of 1 when a country experiences a systemic banking, currency, and debt crisis – is constructed according to the list compiled by Laeven and Valencia (2008), and expanded by recent episodes outlined by Baldacci et al. (2011). The binary variable for IMF-supported stabilization programs is based on lending arrangements,[8] while the binary variable for fiscal rules is drawn from the IMF's Fiscal Rules Dataset.[9]

[6] A detailed description of data sources is presented in the Data Appendix.

[7] While this database (available at http://www.imf.org/external/np/fad/histdb/index.htm) covers an unbalanced panel of 55 countries (24 advanced and 31 emerging-market economies) over the period 1800–2012, we use a balanced panel of 49 countries, excluding Bulgaria, Haiti, Iran, Nicaragua, Romania, and Russia from the original database due to missing observations.

[8] The list of IMF lending arrangements, which are similar to a line of credit and require a country to observe specific terms in order to be eligible to receive a disbursement or to maintain a flexible or precautionary line of credit, is available at http://www.imf.org/external/np/fin/tad/extarr1.aspx.

[9] Schaechter et al. (2012) provide detailed information on the Fiscal Rules Dataset, which is available at http://www.imf.org/external/datamapper/FiscalRules/map/map.htm.

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We test for the stationarity of the variables by applying the Im-Pesaran-Shin and Fisher-type unit root tests for dynamic heterogonous panels.[10] The results, presented in Appendix Table A4, indicate significant test statistics to reject the presence of a unit root in the panel dataset. Additionally, we conduct a test for slope homogeneity using a bootstrapped Hausman test of poolability across advanced and developing countries, and find that the slope coefficients of country groups are similar so that poolability cannot be rejected.[11]

4. EMPIRICAL MODEL AND STRATEGY

Following the existing literature, we model fiscal reaction functions using a range of potential determinants. We build on the model-based fiscal sustainability approach developed by Bohn (1998) and expanded by Fatás and Mihov (2003), Galí and Perotti (2003), and Alesina and Tabellini (2008). In our dynamic panel context, the estimated equation takes the following form:

$$CAPB_{i,t} = \lambda_t + \theta CAPB_{i,t} + \mu GD_{i,t-1} + \rho OG_{i,t-1} + \beta X_{i,t} + \gamma Z_{i,t} + \varepsilon_{i,t} \qquad (1)$$

where $CAPB_{i,t}$ is the cyclically-adjusted primary budget balance scaled by potential GDP in country i at time t, λ_i is a country-specific intercept (fixed effect) accounting for heterogeneity, $CAPB_{i,t-1}$ is the lagged cyclically-adjusted primary balance, $GD_{i,t-1}$ is gross government debt as a share of GDP in country i at time t–1,[12] $OG_{i,t\text{-}1}$ is the lagged output gap,[13] $X_{i,t}$ represents a vector of macro-financial variables for country i at time t, including real GDP per capita, consumer price inflation, interest rates, real exchange rate, domestic credit, stock market capitalization, residential property prices, natural resource rents, and trade openness. $Z_{i,t}$ denotes a matrix of demographic, institutional, and political variables, including population, old-age dependency ratio, a composite index of political regime type, measures of government fragmentation, bureaucratic quality, and corruption, and binary variables for crisis episodes, elections, fiscal rules and IMF programs. $\varepsilon_{i,t}$ is the error term.

We also estimate the determinants of fiscal policy volatility, as measured by the standard deviation of the cyclically-adjusted primary budget balance. Using the above-outlined model, we estimate the equation for the volatility of fiscal reaction functions in the following form:

$$V\_CAPB_{i,t} = \lambda_t + \theta V\_CAPB_{i,t-1} + \mu GD_{i,t-1} + \rho OG_{i,t-1} + \beta Y_{i,t} + \gamma Z_{i,t} + \varepsilon_{i,t} \qquad (2)$$

where $V\_CAPB_{i,t}$ is the standard deviation of the cyclically-adjusted primary balance in country i at time t, $V\_CAPB_{i,t-1}$ is the lagged standard deviation of the cyclically-adjusted primary balance, $GD_{i,t-1}$ is gross government debt as a share of GDP in country i at time t–1, $OG_{i,t-1}$ is the lagged output gap, $Y_{i,t}$ represents a vector of economic and financial variables for country i at time t, including real GDP per capita, the standard deviation of real GDP growth, the level and standard deviation of inflation, the level and standard deviation of interest rates, the level and standard

[10] Descriptive statistics for all variables are presented in Appendix Table A3. In Appendix Table A4, we report results from the Im-Pesaran-Shin (2003) test and Fisher-type tests using ADF and PP tests for unbalanced panels. The Im-Pesaran-Shin test has good small sample performance, while the Fisher-type test uses *p*-values obtained by Monte Carlo simulations from unit root tests for each cross-section. Unlike the Im-Pesaran-Shin test, the Fisher-type tests do not require a balanced panel.

[11] Conventional tests for slope homogeneity – such as a Chow test and the Roy-Zellner Wald-type χ^2 – are less accurate and tend to reject poolability too often even when is true. Bun (2004) tests poolability on dynamic regressions and finds that the classical asymptotic tests tend to over-reject poolability, while bootstrap method tests are more accurate (see also Baltagi, 2008 and Di Iorio and Fachin, 2012).

[12] We also test the square and cubic function of government debt with the aim of capturing nonlinear effects of debt accumulation beyond a certain threshold, but find it to be an insignificant factor in the regressions.

[13] Some studies include the output gap at time *t* as a regressor, but we prefer the lagged output gap, as policymakers may react to past conditions. Moreover, there could be measurement errors in real time, which in turn suggests that data and forecast revisions influence fiscal policy behavior (Beetsma and Giuliodori, 2008).

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deviation of real exchange rate, domestic credit, the level and standard deviation of stock market capitalization, the level and standard deviation of residential property prices, natural resource rents, and trade openness. $Z_{i,t}$ denotes the same set of demographic, institutional, and political variables as defined in Equation 1.

We estimate these models using the system GMM estimator, which corrects for biases associated with endogenous regressors and country-specific heterogeneity. Econometric complications may emerge using the standard estimators, such as the pooled ordinary least squares (OLS) method, because the lagged dependent variable is correlated with the error term, even if we assume that the disturbances are themselves not autocorrelated. One possible solution is the system GMM estimator developed by Arellano and Bover (1995) and Blundell and Bond (1998), which corrects for potential biases associated with endogenous regressors and the persistence of the dependent variable. Although the two-step system GMM estimator is superior in estimating regression models with instrumental variables, it is less reliable in small samples and systematically underestimates the real standard deviation of the estimates compared to the one-step system GMM estimator (Blundell, Bond, and Windmeijer, 2000). Therefore, we prefer the one-step approach in our benchmark estimations and present empirical findings based on the two-step estimator as a robustness check.[14]

5. INTERPRETING EMPIRICAL RESULTS

The empirical findings on the determinants of fiscal reaction functions and volatility of fiscal policy are discussed below in Section A and B, respectively. We estimate a standard model of fiscal sustainability – relating the cyclically-adjusted primary budget balance to its lagged value, lagged government debt, and the lagged output gap – and present the complete set of the results in Table 1.[15] In Tables 2–4, we present our findings based on the system GMM estimation methodology and using a list of potential determinants of fiscal reaction functions.[16] Using a balanced panel, we present two different versions of our benchmark specifications estimated with the one-step system GMM approach in the first and second columns of Tables 2–4 and with the two-step system GMM estimator in the sixth and seventh columns. In Table 5, we present the estimation results for the determinants of fiscal policy volatility, following the same multivariate panel regression approach using one- and two-step system GMM estimators.

5.1. Determinants of Fiscal Policy Behavior

Discretionary fiscal policy has a considerable degree of persistence both in advanced and emerging-market economies. The lagged cyclically-adjusted primary budget balance has a positive and statistically significant coefficient across all specifications of the model, as presented in Tables 1 and 2. The extent of policy inertia is also evident when we estimate fiscal reaction functions separately for advanced and emerging market economies during the period 1990–2012, as presented in Table 3 and Table 4, respectively. Although there are small differences between group coefficients relative to their standard errors, the lagged cyclically adjusted primary budget balance appears to have a greater effect on fiscal policy behavior in advanced economies (0.8–0.9) than in emerging market economies (0.6–0.7).

[14] With the two-step system GMM model, we test the robustness to the reduction of instruments by collapsing the instrument set as suggested by Roodman (2009) and implementing a small sample correction procedure recommended by Windmeijer (2005).

[15] Reduced-form fiscal reaction functions are estimated using Generalized Least Squares (GLS) and one- and two-step system GMM estimators.

[16] All specifications are based on a balanced panel, with the exception of specifications in the third and fifth columns of Tables 2–4. Our panel becomes unbalanced only when we include long-term bond yields and residential property prices, which are not significant factors in either column.

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The fiscal policy stance takes into account the government's intertemporal budget constraint and, therefore, long-run solvency concerns. According to Bohn (1998), the coefficient of the debt variable in the model of fiscal reaction functions must be greater than zero to ensure the sustainability of government finances. As predicted by the theoretical model, we find that the coefficient on public debt is positive and statistically significant in most specifications. This positive response of the cyclically-adjusted primary budget balance to a higher stock of public debt is a robust indication of a pattern of fiscal policy behavior that takes into account the government's intertemporal budget constraint and, therefore, long-run fiscal solvency concerns. Furthermore, we observe a similar pattern of behavior when we estimate fiscal reaction functions separately for advanced and emerging market economies, although the fiscal policy response to the level of public debt is stronger in advanced economies than in emerging market economies.

Discretionary fiscal policy tends to be procyclical both in advanced and emerging market economies. The lagged output gap has a negative and statistically significant coefficient in most specifications of the model, indicating that the policy stance tends to be procyclical. This empirical result is in contradiction to both the standard Keynesian predictions and the tax-smoothing theory, but aligns with a large body of empirical studies that find a pattern of procyclical behavior. Contrary to most of the previous literature, however, we find that the estimated coefficient on the lagged output gap tends to be of a larger magnitude in advanced economies, compared to emerging-market economies, which indicates that the degree of procyclicality is greater in advanced economies.

The inclusion of the contemporaneous output gap, instead of the lagged output gap, points towards countercyclical behavior. This finding is consistent with studies that use real-time data instead of expost observations.[17] Following the approach implemented by Bernoth, Hallet, and Lewis (2008) and Cimadomo (2012), we take the previous year's projections (t–1) for the output gap and cyclically-adjusted primary budget balance as reported in years (t) in the December issues of the OECD Economic Outlook, and estimate our model for 19 OECD countries (which form a subset of our full panel). The coefficients remain similar in size and significance, but the results indicate that the coefficient on the output gap becomes positive when fiscal policy reaction functions are estimated with real-time data. One explanation is that OECD countries tend to plan a countercyclical fiscal strategy, which can turn out to be procyclical in implementation for a variety of reasons including forecast errors, delays in implementation, and policy divergence.

[17] An emerging strand of the empirical literature on fiscal reaction functions, using *real time* data instead of *ex post* observations, finds countercyclical fiscal behavior, especially in advanced economies (see, for example, Forni and Momigliano, 2004; Golinelli and Momigliano, 2006; Cimadomo, 2007; Bernoth, Hallet, and Lewis, 2008). Although *real time* data may yield empirically better performing descriptions of fiscal policy behavior, such figures are available only for a limited number of mostly advanced economies.

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Table 1
Fiscal Reaction Functions

Dependent variable: Cyclically adjusted primary budget balance (CAPB)																		
Panel Specification[4)]	**All Countries**						**Advanced**						**Developing**					
	GLS[1)]	**GLS[1)]**	**SYS GMM[2)]**	**SYS GMM[2)]**	**2S-SYS GMM[3)]**	**2S-SYS GMM[3)]**	**GLS[1)]**	**GLS[1)]**	**SYS GMM[2)]**	**SYS GMM[2)]**	**2S-SYS GMM[3)]**	**2S-SYS GMM[3)]**	**GLS[1)]**	**GLS[1)]**	**SYS GMM[2)]**	**SYS GMM[2)]**	**2S-SYS GMM[3)]**	**2S-SYS GMM[3)]**
Lagged CAPB	0.77***	0.77***	0.82***	0.83***	0.81***	0.78***	0.83***	0.81***	0.91***	0.90***	0.86***	0.90***	0.66***	0.68***	0.67***	0.68***	0.63***	0.68***
	(0.02)	(0.02)	(0.05)	(0.05)	(0.07)	(0.06)	(0.02)	(0.02)	(0.04)	(0.04)	(0.09)	(0.10)	(0.03)	(0.03)	(0.08)	(0.09)	(0.08)	(0.10)
Lagged debt	**0.01*****	**0.01*****	**0.01****	**0.01****	0.01	-0.01	**0.01*****	**0.01*****	**0.01***	**0.01*****	0.01	0.01	**0.01*****	**0.01*****	0.01	0.01	**0.03***	0.02
	(0.00)	(0.00)	(0.00)	(0.00)	(0.01)	(0.01)	(0.00)	(0.00)	(0.00)	(0.00)	(0.01)	(0.01)	(0.00)	(0.00)	(0.00)	(0.00)	(0.01)	(0.02)
Output gap		-0.07***		-0.09**		-0.06		-0.03		-0.06		-0.02		-0.08***		-0.11**		-0.06
		(0.02)		(0.04)		(0.04)		(0.03)		(0.05)		(0.05)		(0.02)		(0.04)		(0.06)
Lagged output gap	-0.11***		-0.16***		0.01		-0.12***		-0.26***		0.03		**-0.09*****		-0.10**		-0.06	
	(0.02)		(0.03)		(0.07)		(0.03)		(0.04)		(0.09)		(0.02)		(0.04)		(0.08)	
Adjusted R^2	0.76	0.75	-	-	-	-	0.80	0.79	-	-	-	-	0.70	0.70	-	-	-	-
Number of countries	49	49	49	49	49	49	24	24	24	24	24	24	26	25	25	25	25	25
Number of years	23	23	23	23	23	23	23	23	23	23	23	23	23	23	23	23	23	23

Notes:
1) Specifications are estimated with the Generalized Least Squares (GLS) assume no constant, the presence of cross-section heteroskedasticity, country-specific fixed effects are included and error term is assumed to follow an AR (1) process.
2) Specifications are estimated with the system Generalized Method of Moments (GMM) assume three lags robust standard errors in parenthesis are consistent to panel specific heteroscedasticity and autocorrelation.
3) Specifications are estimated with the two-step system Generalized Method of Moments (GMM) assuming collapsed instrument matrix, two lags and finite sample corrected standard errors.
4) Robust standard errors reported in parenthesis ***, ** and * denote significance at 1 percent, 5 percent, and 10 percent levels, respectively.

Source: Authors' estimations.

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Economic development and, to a lesser extent, consumer price inflation influence fiscal reaction functions. The coefficient on per capita income indicates that the cyclically-adjusted primary balance increases, on average, by about 0.3–0.4 percentage points with each percentage point increase in the level of real GDP per capita. This finding is consistent with previous empirical studies that estimate income elasticity in an interval of 0.3–0.6, and suggests a more procyclical fiscal reaction. Furthermore, the results indicate that the positive coefficient on per capita income is of a marginally larger magnitude in advanced economies than in emerging market economies. The magnitude of the estimated coefficient on inflation varies with specifications, and remains weakly significant. In particular, the impact of inflation on discretionary fiscal policy becomes less robust when we estimate the model separately for advanced and emerging-market economies. While the coefficient remains positive and mostly significant for advanced economies, it turns out to be negative, but statistically insignificant, in one of the specifications for emerging-market economies.

Financial developments appear to exert a significant effect on discretionary fiscal policy in advanced and emerging-market economies. The extent of financial development, measured by the ratio of domestic credit to GDP, tends to have two opposing effects on the cyclically-adjusted primary balance: enabling deficit financing while imposing market discipline on fiscal policy behavior. In this context, short-term interest rates appear to have a statistically significant influence, but their economic magnitude is negligible across various specifications. The impact of short-term interest rates is significantly greater in advanced economies than in developing countries, suggesting stronger linkages between monetary and fiscal policymaking in advanced economies. Long-term bond yields, on the other hand, do not appear to be an empirically important factor when we estimate the model for the full set of countries. However, our dataset becomes unbalanced when we introduce long-term bond yields into the model, because of limited data availability in emerging-market economies. Indeed, when the model is estimated for the sample of advanced countries, long-term bond yields turn out to have a positive, but still statistically insignificant effect, which may reflect market discipline (i.e. fiscal consolidation efforts when the cost of debt service increases). Likewise, we find that stock market capitalization and property prices have a positive and statistically significant effect on fiscal policy behavior across all countries, albeit the cyclically-adjusted primary budget balance is not corrected for asset price changes.[18]

Structural features of the economy tend to be an important determinant of discretionary fiscal policy. While trade openness appears to have no significant effect, change in terms of trade turns out to be an important determinant of the cyclically-adjusted primary budget balance, particularly in emerging-market economies.[19] We also find that resource dependence – measured by natural resource rents as a share of GDP – tends to have a statistically significant effect on the fiscal policy stance. This result, however, should not be taken as an indication of prudent fiscal behavior in natural resource–rich countries. Indeed, our dependent variable – the cyclically-adjusted primary budget balance – may not appropriately measure the fiscal stance in economies, where commodity price cycles tend to have a significant bearing on fiscal outcomes.[20] Moreover, the findings indicate that real exchange rate appreciation has a statistically significant negative effect on fiscal behavior, which possibly reflects the fall in revenues in local currency terms

[18] Although the effect of residential property prices turns out to be negative in our sample of developing countries, this could be a dubious result because of limited data availability.

[19] The growth literature shows that terms-of-trade shocks are a key determinant of macroeconomic performance in developing countries (see, for example, Cashin and Pattillo, 2000). Indeed, the estimated coefficient on trade openness is negative in the case of emerging-market economies, which may be a reflection of a higher degree of vulnerability to external shocks.

[20] A better measure of the fiscal stance in resource–rich countries is the cyclically-adjusted primary balance excluding resource-based revenues, but these figures are not readily available for our sample of countries.

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when the exchange rate appreciates. Finally, the old-age-dependency ratio appears to influence fiscal policy behavior, with a reasonably negative and statistically significant coefficient. Its economic magnitude is greater among advanced economies, which tend to have a less favorable demographic profile.

Institutional and political factors appear to influence fiscal behavior, varying in direction and degrees of magnitude. Government fragmentation has no strong effect on fiscal policy in advanced economies, but becomes empirically significant in emerging-market economies. On the other hand, political regime type – measured by a composite index ranging from autocracy to democracy – is not a statistically significant factor in determining fiscal policy behavior in advanced and emerging market economies. The impact of electoral cycles on discretionary fiscal policy decisions, as measured by a binary variable that takes the value of 1 in the year preceding elections to capture opportunistic behavior, does not appear to be statistically significant.[21] We also find that the quality of bureaucratic institutions has a positive and statistically significant effect on discretionary fiscal policy, suggesting that better institutions are associated with a more prudent structural fiscal position. Likewise, corruption – measured by an index ranging from 0 (most corrupt) to 6 (least corrupt) – turns out to be a statistically significant factor. That is, a lower degree of corruption improves the cyclically-adjusted primary balance. Furthermore, we find that the interaction of government fragmentation and corruption is a significant factor, indicating a higher degree of vulnerability to corruption in a politically-fragmented environment.

Crisis episodes have a large and statistically significant effect on fiscal policy behavior across all countries. The fiscal impact of crisis episodes, as measured by a dummy variable for banking, currency and debt crises, appears to be larger in advanced economies, but is still substantial in emerging market economies. Furthermore, fiscal reaction functions tend to differ during periods of crisis compared to pre-crisis periods. Using the full set of advanced and emerging-market economies, we compare three different subperiods – the 1990s, 2000–07 and 2008–12 – and present the results in Appendix Table A5.[22] The findings indicate that the degree of persistence in discretionary fiscal policy (as measured by the coefficient on the lagged cyclically-adjusted primary balance) is significantly greater in the post-crisis period. While turning even more procyclical after the 2008–09 crisis, fiscal policy behavior appears to have become less responsive to the government's intertemporal budget constraint and, therefore, long-run fiscal solvency concerns. We also check the interaction of crisis episodes and government debt and find a negative coefficient, as expected. Even though it is not empirically significant, this result implies that crisis episodes have a more pronounced fiscal impact in countries with a higher level of indebtedness.

IMF-supported stabilization programs tend to have a significant effect on fiscal behavior, while fiscal rules turn out to play an insignificant role. IMF-supported stabilization programs have a statistically significant effect on fiscal policy behavior across various specifications of the model. This finding remains valid when we estimate the model separately for subsamples of advanced and emerging-market economies. Fiscal rules, on the other hand, come out to be statistically insignificant in our broad set of countries. This may, however, reflect the fact that most fiscal rules in our sample are not defined in terms of the cyclically-adjusted primary balance, which is our measure of fiscal policy behavior.

21 Although the literature tends to focus on pre-electoral expansionary fiscal policies, the results do not change when we include the election dummy without a lag.

22 Additional tests are performed for parameter consistency and homogeneity during three subperiods.

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Table 2
Determinants of Fiscal Reaction Functions – All Countries

Dependent variable: Cyclically adjusted primary budget balance (*t*)							
	one-step GMM estimation					**two-step GMM estimation**	
	1	**2**	**3**	**4**	**5**	**6**	**7**
Cyclically adjusted primary balance (*t*–1)	0.816***	0.787***	0.792***	0.810***	0.853***	0.805***	0.728***
	[0.046]	[0.050]	[0.062]	[0.050]	[0.052]	[0.067]	[0.054]
Gross government debt (*t*–1)	**0.003***	**0.018***	**0.020***	0.006	**0.025****	0.008	0.008
	[0.002]	[0.010]	[0.011]	[0.005]	[0.010]	[0.008]	[0.035]
Output gap (*t*–1)	**-0.163****	**-0.119*****	**-0.113*****	**-0.143****	**-0.188*****	0.008	0.047
	[0.031]	[0.036]	[0.032]	[0.033]	[0.061]	[0.072]	[0.091]
Real GDP per capita (log)		0.261**					0.435**
		[0.122]					[0.189]
Inflation			0.030		0.141**		
			[0.019]		[0.066]		
Short-term interest rate		0.001**		-0.001*			0.001***
		[0.000]		[0.000]			[0.000]
Real exchange rate		-0.006	0.001	-0.006*			-0.015***
		[0.004]	[0.004]	[0.003]			[0.005]
Change in terms of trade				0.023*			
				[0.012]			
Domestic credit (log)		-0.613***	-0.347*		-0.908***		-0.624
		[0.215]	[0.202]		[0.314]		[0.430]
Stock market capitalization		0.008***	0.007***		0.006**		0.007*
		[0.002]	[0.003]		[0.003]		[0.004]
Population (log)			-0.146**				
			[0.068]				
Corruption			0.240***				
			[0.077]				
Bureaucratic quality					0.337*		
					[0.198]		
Crisis			-0.966**		-1.305**		
			[0.421]		[0.528]		
IMF-supported program		0.573*		0.642**			0.767
		[0.285]		[0.242]			[0.714]
Number of observations	1171	1073	874	1078	565	1171	1073
Number of countries	49	48	43	49	32	49	48
Specification tests (p-values)							
Arellano-Bond AR(1) test	0.001	0.001	0.005	0.002	0.046	0.000	0.001
Arellano-Bond AR(2) test	0.753	0.893	0.923	0.907	0.249	0.773	0.916
Hansen-Sargent test	0.999	0.999	0.999	0.999	0.999	0.060	0.054

Notes:

1) The table reports one- and two-step system GMM dynamic panel estimations carried out using the 'xtabond2' package in Stata. The one-step estimation uses three lags with robust standard errors in brackets, consistent to panel specific heteroscedasticity and autocorrelation. The two-step estimation uses collapsed instrument matrix, two lags and finite sample corrected standard errors in brackets. The AR(1) and AR(2) tests report the p-values for the first and second order residual autocorrelation in the first differenced equation, providing no evidence for significant second order autocorrelation. Hansen-Sargent test for overindentifying restrictions provides the probability value for H_0 joint validity of the instruments. Higher probability value suggests that the instruments are exogenous and not correlated with the error term. The test is robust but grows weaker with higher number of moment conditions.

2) All results in the table are presented in easily comparable format. Regressions 1 and 2 (one step GMM) have exactly the same specification as regressions 6 and 7 (two steps GMM).

3) Regressions include the following control variables – long term bond yield, trade openness, property prices, natural resource rents, old age dependency ratio, regime*fragmentation, fragmentation*corruption, lagged election and fiscal rule, which were found to be insignificant.

4) *, **, *** indicates significance at 10 percent ($p < 0.10$), 5 percent ($p < 0.05$) and 1 percent ($p < 0.01$), respectively.

Source: Authors' estimations.

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Table 3
Determinants of Fiscal Reaction Functions – Advanced Countries

Dependent variable: Cyclically adjusted primary budget balance (*t*)							
	one-step GMM estimation					two-step GMM estimation	
	1	2	3	4	5	6	7
Cyclically adjusted primary balance (*t*–1)	0.909***	0.855***	0.900***	0.886***	0.824***	0.855***	0.826***
	[0.036]	[0.070]	[0.041]	[0.054]	[0.066]	[0.089]	[0.083]
Gross government debt (*t*–1)	**0.003***	**0.018****	**0.024*****	**0.010***	**0.021****	0.004	0.049
	[0.002]	[0.008]	[0.008]	[0.006]	[0.009]	[0.004]	[0.033]
Output gap (*t*–1)	-0.262***	-0.178***	-0.223***	-0.218***	**-0.242*****	0.033	0.085
	[0.039]	[0.046]	[0.047]	[0.037]	[0.063]	[0.091]	[0.097]
Real GDP per capita (log)		0.222					0.532*
		[0.168]					[0.272]
Inflation			0.163*		0.296**		
			[0.082]		[0.093]		
Short-term inerest rate		0.094***		0.089***			0.070
		[0.030]		[0.022]			[0.063]
Real exchange rate		0.002	0,015	**-0.014*****			-0.014
		[0.014]	[0.010]	[0.004]			[0.022]
Domestic credit (log)		-0.894**	-0.628**		-0.698*		-1.139**
		[0.367]	[0.287]		[0.366]		[0.494]
Stock market capitalization		0.009***	0.007**		0.008**		0.009***
		[0.003]	[0.003]		[0.003]		[0.002]
Natural resource rents		0.172***		0.129***	0.214**		0.266***
		[0.059]		[0.038]	[0.080]		[0.062]
Population (log)			-0.137**				
			[0.067]				
Old age dependency ratio		-0.026	**-0.052***				-0.132
		[0.040]	[0.025]				[0.084]
Crisis			-1.070**		-1.151***		
			[0.550]		[0.593]		
IMF-supported program		3.457*		3.203**			1.530
		[1.734]		[1.500]			[1.505]
Number of observations	552	517	507	506	473	552	517
Number of groups	23	23	23	23	23	23	23
Specification tests (p-values)							
Arellano-Bond AR(1) test	0.047	0.047	0.049	0.058	0.068	0.028	0.030
Arellano-Bond AR(2) test	0.475	0.228	0.321	0.301	0.289	0.450	0.192
Hansen-Sargent test	0.999	0.999	0.999	0.999	0.999	0.580	0.585

Notes:
1) The table reports one- and two-step system GMM dynamic panel estimations carried out using the 'xtabond2' package in Stata. The one-step estimation uses three lags with robust standard errors in brackets, consistent to panel specific heteroscedasticity and autocorrelation. The two-step estimation uses collapsed instrument matrix, two lags and finite sample corrected standard errors in brackets. For all estimations the difference-in-Hansen test reports the p-values for the null hypothesis of additional moment conditions validity. The AR(1) and AR(2) tests report the p-values for the first and second order residual autocorrelation in the first differenced equation, providing no evidence for significant second order autocorrelation. Hansen-Sargent test for overindentifying restrictions provides the probability value for H_0 joint validity of the instruments. Higher probability value suggests that the instruments are exogenous and not correlated with the error term. The test is robust but grows weaker with higher number of moment conditions.
2) All results in the table are presented in easily comparable format. Regressions 1 and 2 (one step GMM) have exactly the same specification as regressions 6 and 7 (two steps GMM).
3) Regressions include the following control variables – long term bond yield, trade openness, chnage in terms of trade, property prices, government fragmentation, regime*fragmentation, fragmentation*corruption, corruption, bureaucratic quality, lagged election and fiscal rule, which were found to be insignificant.
4) *, **, *** indicates significance at 10 percent ($p < 0.10$), 5 percent ($p < 0.05$) and 1 percent ($p < 0.01$), respectively.

Source: Authors' estimations.

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Table 4
Determinants of Fiscal Reaction Functions – Emerging Markets

Dependent variable: Cyclically adjusted primary budget balance (*t*)							
	one-step GMM estimation					**two-step GMM estimation**	
	1	**2**	**3**	**4**	**5**	**6**	**7**
Cyclically adjusted primary balance (*t*–1)	0.674***	0.632***	0.600***	0.649***	0.606***	0.628***	0.662***
	[0.085]	[0.081]	[0.108]	[0.089]	[0.113]	[0.077]	[0.068]
Gross government debt (*t*–1)	0,005	**0.016***	**0.019*****	**0.011***	**0.052*****	**0.027***	**0.047****
	[0.003]	[0.008]	[0.007]	[0.006]	[0.015]	[0.014]	[0.019]
Output gap (*t*–1)	**-0.103****	-0.071	**-0.091***	**-0.071****	0.025	-0.062	0.200
	[0.042]	[0.042]	[0.046]	[0.037]	[0.086]	[0.076]	[0.075]
Real GDP per capita (log)		0.357***					0.423
		[0.125]					[0.210]
Inflation			0.038**		-0.058		
			[0.015]		[0.060]		
Short-term interest rate		0.000**		-0.001**			0.001**
		[0.000]		[0.000]			[0.000]
Real exchange rate		-0.010**	-0.006**	-0.006***			-0.008
		[0.005]	[0.004]	[0.002]			[0.006]
Change in terms of trade				0.024**			
				[0.009]			
Domestic credit (log)		**-0.572*****	-0.077		0.304		-0.844**
		[0.187]	[0.182]		[0.417]		[0.380]
Stock market capitalization		0.106**	0.011**		0.006**		0.018***
		[0.004]	[0.005]		[0.002]		[0.006]
Property prices (log)					-0.594**		
					[0.182]		
Natural resource rents		0,023		**0.029****	-0.017		0.039*
		[0.156]		[0.013]	[0.033]		[0.022]
Old age dependency ratio		**-0.065****	0.002				-0.048
		[0.026]	[0.025]				[0.039]
Government fragmentation					1.275**		
					[0.559]		
Crisis			-0.085		-2.136***		
			[0.414]		[0.467]		
IMF-supported program		0.382*		0.415*			0.024
		[0.222]		[0.206]			[0.344]
Number of observations	619	556	367	572	92	619	556
Number of groups	26	25	20	26	9	26	25
Specification tests (p-values)							
Arellano-Bond AR(1) test	0.002	0.004	0.007	0.006	0.069	0.003	0.006
Arellano-Bond AR(2) test	0.457	0.482	0.498	0.489	0.699	0.466	0.530
Hansen-Sargent test	0.999	0.999	0.999	0.999	0.999	0.212	0.161

Notes:
1) The table reports one- and two-step system GMM dynamic panel estimations carried out using the 'xtabond2' package in Stata. The one-step estimation uses three lags with robust standard errors in brackets, consistent to panel specific heteroscedasticity and autocorrelation. The two-step estimation uses collapsed instrument matrix, two lags and finite sample corrected standard errors in brackets. For all estimations the difference-in-Hansen test reports the p-values for the null hypothesis of additional moment conditions validity. The AR(1) and AR(2) tests report the p-values for the first and second order residual autocorrelation in the first differenced equation, providing no evidence for significant second order autocorrelation. Hansen-Sargent test for overindentifying restrictions provides the probability value for H_0 joint validity of the instruments. Higher probability value suggests that the instruments are exogenous and not correlated with the error term. The test is robust but grows weaker with higher number of moment conditions.
2) All results in the table are presented in easily comparable format. Regressions 1 and 2 (one step GMM) have exactly the same specification as regressions 6 and 7 (two steps GMM).
3) Regressions include the following control variables – long term bond yield, trade openness, population, regime*fragmentation, fragmentation*corruption, corruption, bureaucratic quality, lagged election and fiscal rule, which were found to be insignificant.
4) *, **, *** indicates significance at 10 percent ($p < 0.10$), 5 percent ($p < 0.05$) and 1 percent ($p < 0.01$), respectively.

Source: Authors' estimations.

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Table 5
Determinants of Fiscal Policy Volatility – All Countries

Dependent variable: Standard Deviation (SD) of the cyclically adjusted primary budget balance (*t*)				
	one-step GMM estimation			two-step GMM estimation
	1	2	3	4
SD of cyclically adjusted primary balance (*t*–1)	0.292***	0.228***	0.251**	0.171***
	[0.060]	[0.071]	[0.088]	[0.056]
Gross government debt (*t*–1)	**0.016*****	**0.017*****	**0.021*****	0.020
	[0.007]	[0.004]	[0.007]	[0.016]
Output gap (*t*–1)	0.081***	0.084***	0.093***	0.024
	[0.018]	[0.015]	[0.020]	[0.030]
SD of real GDP growth		0.011	**0.051***	
		[0.028]	[0.021]	
SD of real exchange rate	**0.020****	0.014		0.023***
	[0.008]	[0.010]		[0.008]
Natural resource rents	0.041***	0.040***		0.039**
	[0.012]	[0.015]		[0.015]
Bureaucratic quality			-0.267**	
			[0.113]	
Fiscal rule		-0.442***		
		[0.161]		
Number of observations/countries	1025/48	1025/48	1076/49	1025/48
Number of instruments	129	131	129	11
Specification tests (p-values)				
Arellano-Bond AR(1) test	0.000	0.000	0.000	0.000
Arellano-Bond AR(2) test	0.869	0.930	0.569	0.796
Hansen-Sargent test	0.999	0.999	0.999	0.106

Notes:
1) The table reports one and two-step System GMM dynamic panel estimations carried out using the ‘xtabond2’ package in Stata (Roodman, 2009). The one-step System GMM uses three lags with robust standard errors in brackets, consistent to panel specific heteroscedasticity and autocorrelation. The two-step System GMM model uses collapsed instrument matrix, two lags and only one instrument for each variable. Windmeijer (2005) finite sample corrected standard errors in brackets are employed. The AR(1) and AR(2) tests report the p-values for the first and second order residual autocorrelation in the first differenced equation, providing no evidence for significant second order autocorrelation. Hansen-Sargent test for overindentifying restrictions provides the probability value for H_0 joint validity of the instruments. Higher probability value suggests that the instruments are exogenous and not correlated with the error term. The test is robust but grows weaker with higher number of moment conditions.
2) All results in the table are presented in easily comparable format. Regression 1 (one-step GMM) have exactly the same specification as regression 4 (two-steps GMM).
3) Regressions include the following control variables – real GDP per capita, standard deviation of cpi, trade openness, domestic credit, standard deviation of stock market cap., government fragmentation, corruption, and crisis, which were found to be insignificant.
4) *, **, *** indicates significance at 10% ($P < 0.10$), 5% ($P < 0.05$) and 1% ($P < 0.01$) respectively.

Source: Authors’ estimations.

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Table 6
Determinants of Fiscal Policy Volatility – Advanced Countries

Dependent variable: Standard Deviation (SD) of the cyclically adjusted primary budget balance (t)				
	one-step GMM estimation			**two-step GMM estimation**
	1	**2**	**3**	**4**
SD of cyclically adjusted primary balance (*t*–1)	0.253***	0.217***	0.214***	0.218**
	[0.066]	[0.074]	[0.056]	[0.091]
Gross government debt (*t*–1)	0.007	**0.009****	**0.010****	0.018
	[0.004]	[0.004]	[0.004]	[0.025]
Output gap (*t*–1)	**0.061****	**0.054****	**0.075*****	0.030
	[0.029]	[0.024]	[0.022]	[0.055]
SD of real GDP growth		0.084*	0.089**	
		[0.049]	[0.043]	
Natural resource rents	0.056***	0.057***		0.073**
	[0.018]	[0.018]		[0.042]
Number of observations/countries	494/23	494/23	506/23	494/23
Number of instruments	129	131	129	11
Specification tests (p-values)				
Arellano-Bond AR(1) test	0.001	0.001	0.000	0.001
Arellano-Bond AR(2) test	0.466	0.455	0.483	0.474
Hansen-Sargent test	0.999	0.999	0.999	0.058

Notes:
1) The table reports one and two-step System GMM dynamic panel estimations carried out using the 'xtabond2' package in Stata (Roodman, 2009). The one-step System GMM uses three lags with robust standard errors in brackets, consistent to panel specific heteroscedasticity and autocorrelation. The two-step System GMM model uses collapsed instrument matrix, two lags and only one instrument for each variable. Windmeijer (2005) finite sample corrected standard errors in brackets are employed. The AR(1) and AR(2) tests report the p-values for the first and second order residual autocorrelation in the first differenced equation, providing no evidence for significant second order autocorrelation. Hansen-Sargent test for overindentifying restrictions provides the probability value for H_0 joint validity of the instruments. Higher probability value suggests that the instruments are exogenous and not correlated with the error term. The test is robust but grows weaker with higher number of moment conditions.
2) All results in the table are presented in easily comparable format. Regression 1 (one-step GMM) have exactly the same specification as regression 4 (two-steps GMM).
3) Regressions include the following control variables - real GDP per capita, standard deviation of cpi,standard deviation of real exchange rate, trade openness, domestic credit, standard deviation of stock market cap., government fragmentation, corruption, bureaucratic quality, crisis and fiscal rule, which were found to be insignificant.
4) *, **, *** indicates significance at 10% (P < 0.10), 5% (P < 0.05) and 1% (P < 0.01) respectively.

Source: Authors' estimations.

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Table 7
Determinants of Fiscal Policy Volatility – Emerging Markets

Dependent variable: Standard Deviation (SD) of the cyclically adjusted primary budget balance (t)				
	one-step GMM estimation			**two-step GMM estimation**
	1	**2**	**3**	**4**
SD of cyclically adjusted primary balance (*t*–1)	0.266***	0.241**	0.293***	0.144**
	[0.083]	[0.090]	[0.125]	[0.057]
Gross government debt (*t*–1)	0.016**	0.014***	0.016**	0.020**
	[0.007]	[0.005]	[0.007]	[0.008]
Output gap (*t*–1)	**0.080****	**0.072*****	**0.080*****	0.026
	[0.025]	[0.019]	[0.022]	[0.030]
SD of real exchange rate	0.019**	0.020*		0.023***
	[0.009]	[0.010]		[0.007]
SD of stock market capitalization	**0.013****	0.009		0.013*
	[0.006]	[0.006]		[0.008]
Natural resource rents	0.033***	0.037**		0.067
	[0.013]	[0.016]		[0.034]
Number of observations/countries	531/25	531/25	570/26	531/25
Number of instruments	129	131	129	11
Specification tests (p-values)				
Arellano-Bond AR(1) test	0.004	0.004	0.002	0.011
Arellano-Bond AR(2) test	0.238	0.278	0.340	0.444
Hansen-Sargent test	0.999	0.999	0.999	0.675

Notes:
1) The table reports one and two-step System GMM dynamic panel estimations carried out using the ‘xtabond2’ package in Stata (Roodman, 2009). The one-step System GMM uses three lags with robust standard errors in brackets, consistent to panel specific heteroscedasticity and autocorrelation. The two-step System GMM model uses collapsed instrument matrix, two lags and only one instrument for each variable. Windmeijer (2005) finite sample corrected standard errors in brackets are employed. The AR(1) and AR(2) tests report the p-values for the first and second order residual autocorrelation in the first differenced equation, providing no evidence for significant second order autocorrelation. Hansen-Sargent test for overindentifying restrictions provides the probability value for H_0 joint validity of the instruments. Higher probability value suggests that the instruments are exogenous and not correlated with the error term. The test is robust but grows weaker with higher number of moment conditions.
2) All results in the table are presented in easily comparable format. Regression 1 (one-step GMM) have exactly the same specification as regression 4 (two-steps GMM).
3) Regressions include the following control variables - real GDP per capita,standard deviation of real GDP, standard deviation of cpi, trade openness, domestic credit, government fragmentation, corruption, bureaucratic quality, crisis and fiscal rule, which were found to be insignificant.
4) *, **, *** indicates significance at 10% ($P < 0.10$), 5% ($P < 0.05$) and 1% ($P < 0.01$) respectively.

Source: Authors’ estimations.

5.2. Determinants of Fiscal Policy Volatility

Higher public debt leads to more volatile fiscal behavior, while fiscal rules and higher institutional quality reduce the volatility of fiscal policy. As presented in Tables 5–7, the lagged volatility of the cyclically-adjusted primary balance has a positive and empirically significant coefficient across all specifications of the model, indicating a high degree of persistence in the volatility of fiscal policy reaction functions over time. The extent of policy volatility appears to be

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higher in emerging market economies, but still economically substantial in advanced economies as well. The level of public debt and the lagged output gap have statistically significant and positive coefficients, indicating that higher government debt or output gap (i.e. above-potential growth) lead to more volatile policy behavior. Real GDP growth volatility has a positive and empirically significant coefficient, particularly in the case of advanced economies, while the volatility of consumer price inflation appears to be insignificant across all countries. Fiscal policy volatility is higher in the presence of natural resource rents, real exchange rate volatility, and stock market volatility, especially in emerging-market economies. On the other hand, we find that fiscal rules and higher institutional quality reduce volatility of fiscal policy, while crisis episodes do not appear to have a statistically significant effect on the volatility of fiscal policy behavior.[23]

The empirical results remain robust when we use an alternative measure of discretionary fiscal policy volatility. Following Fatás and Mihov (2003), we introduce an alternative regression-based measure of average volatility of policy changes using the standard deviation of the residuals from country-specific regressions (Appendix Table A6). The residuals quantitatively estimate discretionary fiscal policy and are drawn from the estimated fiscal reaction function of the cyclically-adjusted fiscal balance on its lagged value, lagged government debt, and lagged output gap. We find that the lagged volatility of discretionary fiscal policy has a positive and significant effect on the cyclically-adjusted fiscal balance volatility, supporting the results in Table 5. These results are also robust to the subsamples of advanced and developing countries as the volatility of discretionary fiscal policy is, with somewhat smaller magnitude, in advanced economies.

6. ROBUSTNESS CHECKS

We perform several specification tests in order to ensure the validity of the system GMM estimations. Firstly, we estimate alternative specifications including additional determinants and arrive at results similar to the baseline models, presented in the first and second columns of Tables 2–4. Secondly, since the one-step system GMM estimator becomes weaker when the number of instruments increases, we test the second and third lags of the endogenous variable to avoid the problem of invalid instruments (correlated with the error term) or weak instruments (only weakly correlated with explanatory variables). Thirdly, we test the robustness of our benchmark one-step system GMM estimator by comparing three different time periods – the 1990s, 2000–08 and 2009–12 (Appendix Table A5). Fourthly, we consider a two-step system GMM estimator, as presented in the last two columns of Tables 2–4, to test the sensitivity of our empirical findings to different GMM estimators.

The two-step system GMM is defined as an alternative model, testing a reduced number of instruments to avoid overidentification while ensuring validity. The two-step estimator is more efficient but can be biased downwards for finite sample inference. We follow Roodman (2009) and estimate the baseline model with two lags and a collapsed instrument matrix, which specifies an instrument for each variable and reduces the size of the instrument matrix by a smaller set of moment conditions. The covariance matrix is robust to the panel-specific autocorrelation and heteroscedasticity improved with the finite sample corrected standard errors developed by Windmeijer (2005). The two-step estimates in Tables 2–4, appear to be consistent in direction and size with the results from the one-step analysis forming a downward interval limit with coefficient that lies between the two bounds given in the sixth and seventh columns and the first and second columns. Some coefficients are not significant when estimated with the two-step GMM model

[23] This is broadly in line with other studies estimating the impact of fiscal rules on discretionary policymaking. For example, Fatás and Mihov (2003), Woo (2009), and Brzozowski and Siwinska-Gorzelak (2010) find that formal constraints tend to lower the volatility of government spending.

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compared to the one-step GMM suggesting the possibility that too many instruments overfit the model (i.e., instrument proliferation problem) resulting in upward-biased estimates.

The test results are robust across all regressions, indicating that our instruments are valid, but weaken with a higher number of instruments. All model specifications satisfy the test statistics for overidentifying restrictions in the instrumental variables, although the number of instruments may be large relative to the number of groups in the one-step approach. Considering that the validity of the instrument set depends on the error structure, we also report the Arellano and Bond (1991) tests AR(1) and AR(2) with p-values for first and second order autocorrelated disturbances in the first-differenced equation. All in all, the tests provide evidence for high first-order autocorrelation (the AR (1) rejects the null hypothesis of no autocorrelation) but yields no evidence for significant second-order autocorrelation (higher p-value of the AR(2) statistics) suggesting correctly-specified models. Finally, we include the Hansen-Sargan test for overidentifying restrictions to check the joint validity of the instruments. The probability values for the null hypothesis – that is, the instruments are valid – are presented in the last row of each table. Higher probability value suggests that the instruments are exogenous and not correlated with the error term. The test is robust across all regressions, indicating that our instruments are valid, but the test becomes weaker with a higher number of instruments.[24]

7. CONCLUSIONS

This paper investigates the main determinants of discretionary fiscal policy and its volatility in advanced and emerging-market economies. The state of public finances deteriorated significantly in the aftermath of the 2008–09 economic crisis. While these extraordinary developments have generally been blamed on fiscal stimulus packages, the severity of the economic downturn, and the bailout of financial institutions, there are no systematic time-series estimates of fiscal reaction functions for a large panel of advanced and emerging market economies over the period 1990–2012. Accordingly, this paper investigates the main determinants of discretionary fiscal policy, as measured by the cyclically-adjusted primary budget balance, and its time-varying volatility. In view of omitted-variables bias and potential endogeneity problems that plague standard econometric techniques commonly used in the empirical literature, this paper adopts a dynamic panel approach based on the system GMM estimator and uses a balanced panel that yields more robust estimations than unbalanced panels.

The empirical results suggest that discretionary fiscal policy is influenced by a range of macro-financial and institutional factors. In line with the existing literature, we show that discretionary fiscal policy is influenced by policy inertia, the level of public debt, and the output gap in both advanced and emerging market economies, although the fiscal policy response to debt accumulation is stronger in advanced economies than in emerging-market economies. In addition, we find that macro-financial factors (such as real exchange rate, financial development, interest rates, asset prices, and natural resource rents) and demographic and institutional factors (such as the old-age dependency ratio, the quality of institutions, and policy anchors such as fiscal rules and IMF-supported stabilization programs) tend to have a significant influence fiscal behavior. The empirical results indicate that these factors tend to influence discretionary fiscal policy differently in advanced and emerging-market economies. Moreover, we show that the degree of policy inertia is significantly greater in the post-crisis period, with fiscal behavior turning more procyclical and becoming less responsive to the government's intertemporal budget constraint. With regards to the volatility of fiscal policymaking, the results indicate that higher

[24] It should be noted that having a large number of moment conditions can overfit the endogenous variables, weakening Hansen tests of instrument validity (Roodman, 2009).

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government debt leads to more volatile policy behavior, while fiscal rules and higher institutional quality reduce volatility of fiscal policy.

Policymakers should aim for a countercyclical fiscal policy stance that takes into account long-run solvency concerns. The empirical results presented in this paper have a number of critical policy implications, as countries continue to struggle with putting public finances on a growth-enhancing and sustainable path. Firstly, discretionary fiscal policy tends to exhibit a pattern of procyclical behavior both in advanced and emerging market economies, worsening, instead of smoothing out, macro-financial oscillations. Secondly, although the fiscal policy stance appears to take into account the government's intertemporal budget constraint across all countries in our sample, the policy response to the level of public debt is stronger in advanced economies than in emerging market economies. In the aftermath of the 2008–09 crisis, however, fiscal policy behavior has become more procyclical and less responsive to the government's intertemporal budget constraint and, therefore, long-run fiscal solvency concerns, especially in advanced economies. In this context, the results show that policymakers need to take into account financial developments, such as the fluctuations in stock market capitalization and property prices that tend to have a significant effect on fiscal policy behavior. Likewise, improving the quality of bureaucratic institutions is a necessary condition to have an effective fiscal policy framework, especially against the risk of corruption in a politically fragmented environment. Finally, even though fiscal rules do not appear to be an empirically significant determinant of fiscal policy behavior, a rule-based fiscal regime is still found to reduce the volatility of fiscal policy over time.

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APPENDIX

Table A1
Country Sample

Advanced 24	Developing 25
Australia	Argentina
Austria	Bolivia
Belgium	Brazil
Canada	Chile
Denmark	China
Finland	Colombia
France	Costa Rica
Germany	Dominican Republic
Greece	Ghana
Iceland	Honduras
Ireland	Hungary
Israel	India
Italy	Indonesia
Japan	Mexico
South Korea	Pakistan
Netherlands	Panama
New Zealand	Paraguay
Norway	Peru
Portugal	Philippines
Spain	Poland
Sweden	South Africa
Switzerland	Thailand
United Kingdom	Turkey
United States	Uruguay
	Venezuela

Notes:
[1)] Country classification is based on per capita income level, export diversification, and degree of integration into the global financial system, according to the IMF's WEO database.

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Table A2
Descriptions of Variables and Data Sources

Variable	Description	Source
Cyclically adjusted primary budget balance	Budget balance net of interest payments and stripping out the effect of the business cycle in percent of potential GDP	IMF/FAD
Gross government debt	Stock of gross general government debt in percent of GDP	WEO
Real GDP per capita	Ratio of real GDP to total population (log)	WEO
Output gap	Deviation of real GDP from its trend in percent of GDP	IMF
Inflation	Annual percentage change in the consumer price index	WEO
Short-term interest rate	Percent	WEO
Long-term bond yield	Percent	WEO
Real exchange rate	Real effective exchange rate index (2005 = 100)	WDI
Domestic credit	Domestic bank lending in percent of GDP (log)	WEO
Market capitalization	Stock market valuation in percent of GDP	WDI
Property prices	Average residential real estate prices (log)	BIS
Natural resource rents	in percent of GDP	WDI
Trade openness	Ratio of exports amd imports to GDP (percent)	WEO
Change in terms of trade	Annual percentage change in the terms of trade index	WEO
Population	Total population in millions (log)	WEO
Political regime	Type of political regime ranging from -10 (strongly autocratic) to 10 (strongly democratic)	Polity IV
Government fragmentation	The probability that two deputies picked random from among the government parties will be of different parties	DPI
Bureaucratic quality	Index	ICRG
Corruption	Index of corruption perception measuring excessive patronage, nepotism, job reservation, secret party funding, and close ties between politics and business	ICRG
Old-age dependency ratio	Number of people aged 65 or over in percent of working-age population (aged 15–64)	WDI
IMF-supported program	Binary variable (taking the value of one when a country implements an IMF-supported program in a given year)	IMF
Crisis	Binary variable (taking the value of one when a country experiences an episode of banking, currency and/or debt crisis)	IMF
Election	Binary variable (taking the value of one in an election year)	DPI
Fiscal rule	Binary variable (taking the value of one when there is a numerical fiscal rule in effect)	IMF/FAD

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Table A3
Summary Statistics

Variables	Obs	Mean	Std-dev	Min	Max
Cyclically adjusted primary budget balance	1174	1.45	3.68	-26.02	21.44
Gross government debt	1176	57.00	31.00	4.10	235.71
Output gap	1176	-0.07	3.24	-19.08	14.26
Real GDP per capita (log)	1127	8.84	1.34	5.40	10.64
Inflation	1127	25.00	265.25	-1.71	7481.70
Short-term interest rate	1127	26.93	312.73	0.00	9394.29
Long-term bond yield	902	8.80	8.14	0.80	96.88
Real exchange rate	1127	102.93	18.55	37.51	227.20
Domestic credit (log)	1127	4.31	0.70	2.35	5.85
Market capitalization	1073	52.94	49.05	0.17	309.45
Property prices (log)	568	5.06	1.18	3.23	9.70
Natural resource rents	1127	3.35	5.78	0.00	47.88
Trade openness	1127	69.33	34.84	13.75	198.77
Terms of trade	1176	104.2	34.7	28.9	678.8
Population (log)	1176	3.14	1.50	0.23	7.21
Regime type	1127	7.85	3.79	-7.00	10.00
Government fragmentation	1127	0.27	0.28	0.00	0.89
Bureaucratic quality	1127	2.95	1.01	0.00	4.00
Corruption	1127	3.62	1.42	0.00	6.00
Old-age dependency ratio	1127	15.78	7.53	5.35	38.03

Source: Authors' calculations.

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Table A4
Unit Root Tests

Variable[3]	Panel tests for unit root in levels[1]			Panel tests for unit root in differences[1]		
	IPS	ADF – Fisher[2]	PP – Fisher[2]	IPS	ADF – Fisher[2]	PP – Fisher[2]
Cyclically adjusted primary budget balance	-4.18***	164.18***	**110.66**	-18.21***	463.58***	658.58***
Gross government debt	**-1.96**	140.14***	**73.91**	-9.64***	287.45***	305.27***
Real GDP per capita	**6.30**	**65.44**	**51.06**	-10.02***	274.35***	313.85***
Output gap	-2.41***	246.28***	122.19**	-14.74***	372.39***	451.78***
Consumer price index	**3.49**	**128.02**	**91.82**	-7.54***	247.82***	254.90***
Short-term int. rate	-29.13***	729.58***	1003.25***	-69.40***	982.19***	1921.26***
Long-term bond yield	-2.93***	134.17***	168.68***	-19.99***	438.64***	712.40***
Real exchange rate	-2.84**	144.17**	**111.67**	-14.67***	370.98***	500.72***
Domestic credit	**-1.83**	**115.24**	**88.14**	-16.84***	429.06***	623.75***
Broad money supply	**0.55**	**93.30**	**53.61**	-12.60***	344.17***	452.70***
Market capitalization	-5.97***	167.57***	148.31***	-20.56***	517.97***	1645.05***
Property prices	0.70	49.50	**55.35**	-2.91***	91.88***	**59.52**
Natural Resource Rents	-5.92***	191.10***	175.69***	-20.33***	508.47***	1331.37***
Trade openness	-3.32***	140.25**	**110,85**	-18.92***	479.47***	660.92***
Terms of Trade	-4.88***	197.32***	178.89***	-22.48***	582.08***	1364.62***
Population	**2.99**	149.07*	**91.73**	-2.42**	235.77***	252.65***

Notes:
1) Im-Pesaran-Shin (2003) test (IPS) and Fisher tests for unbalanced panels allow for heterogeneous coefficients. All pooled unit root tests include individual intercept and individual linear trend. Four lag length selection based on Schwarz Info Criterion
2) Probabilities for Fisher-type tests using ADF and PP tests are computed with an asymptotic Chi -square distribution. IPS test assumes asymptotic normality.
3) The null hypothesis is that of a unit root process - Ho: all of the time series in the panel are nonstationary; i.e., rejection of the null means that the variables are stationary. The symbols * and ** denote significance at the 5 percent and 1 percent level, respectively.
4) All economic and financial variables are stationary when differenced except property prices. Non stationary variables are in bold. Rejection of the PP – Fisher panel unit root test in differences is likely due to missing values.

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Table A5
System GMM Dynamic Panel by Period – All Countries

Sample Specifications	1990–1999		2000–2007		2008–2012	
	1	**2**	**1**	**2**	**1**	**2**
Cyclically adjusted primary balance (*t*–1)	**0.657*****	**0.573*****	**0.721*****	**0.436*****	**0.865*****	**0.866*****
	[0.139]	[0.139]	[0.070]	[0.099]	[0.050]	[0.059]
Gross government debt (*t*–1)	0.006	0.021	**0.014*****	-0.004	-0,001	0.039
	[0.006]	[0.018]	[0.004]	[0.018]	[0.002]	[0.026]
Output gap (*t*–1)	-0.012	0.038	**-0.100*****	**-0.121***	**-0.300*****	**-0.183****
	[0.040]	[0.058]	[0.035]	[0.061]	[0.069]	[0.089]
Real GDP per capita (log)		**0.478*****		**0.910*****		-0.078
		[0.177]		[0.253]		[0.264]
Short-term interest rate		0.001**		0.019		0.057
		[0.000]		[0.013]		[0.064]
Real exchange rate		**-0.014****		**-0.030*****		0.007
		[0.007]		[0.008]		[0.009]
Domestic credit (log)		**-0.798****		**-0.752***		-0.879
		[0.373]		[0.381]		[0.566]
Stock market capitalization		**0.012*****		0.005		0.012
		[0.004]		[0.004]		[0.007]
Old-age dependency ratio		-0,038		-0.082		-0.009
		[0.040]		[0.063]		[0.067]
Government fragmentation*Corruption		-0.017		**0.410***		0.232
		[0.104]		[0.198]		[0.178]
Number of observations	488	452	392	384	245	237
Number of countries	49	48	49	48	49	48
Specification tests (p-values)						
Arellano-Bond AR(1) test	0.001	0.001	0.001	0.001	0.117	0.112
Arellano-Bond AR(2) test	0.526	0.647	0.929	0.876	0.479	0.448
Hansen-Sargent test	0.430	0.847	0.426	0.730	0.118	0.133

Notes:
1) The table reports one-step system GMM dynamic panel estimations carried out using the 'xtabond2' package in Stata. The one-step estimation uses three lags with robust standard errors in brackets, consistent to panel specific heteroscedasticity and autocorrelation. The two-step estimation uses collapsed instrument matrix, two lags and finite sample corrected standard errors in brackets. The AR(1) and AR(2) tests report the p-values for the first and second order residual autocorrelation in the first differenced equation, providing no evidence for significant second order autocorrelation. Hansen-Sargent test for overindentifying restrictions provides the probability value for H_0 joint validity of the instruments. Higher probability value suggests that the instruments are exogenous and not correlated with the error term. The test is robust but grows weaker with higher number of moment conditions.
2) All results in the table are presented in easily comparable format. Regressions 1 and 2 have exactly the same specification throughout all periods.
3) Regressions include the following control variables - natural resource rents, old age dependency ratio and IMF-supported program, which were found to be insignificant.
4) *, **, *** indicates significance at 10 percent ($p < 0.10$), 5 percent ($p < 0.05$) and 1 percent ($p < 0.01$), respectively.

Source: Authors' estimations.

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Table A6
Residual-based Volatility – All Countries

Dependent variable: Standard Deviation (SD) of the residuals from country-specific regressions[1]				
	one-step GMM estimation			two-step GMM estimation
	1	**2**	**3**	**4**
SD of residuals of cyclically adjusted primary balance (t–1)	**0.310*****	**0.281*****	**0.284*****	**0.066*****
	[0.068]	[0.077]	[0.067]	[0.113]
Gross government debt (t–1)	**0.026*****	**0.027****	**0.028****	0.021
	[0.018]	[0.010]	[0.012]	[0.016]
Output gap (t–1)	**0.096*****	**0.097*****	**0.097*****	0.001
	[0.029]	[0.030]	[0.031]	[0.040]
SD of real GDP growth		0.051*	0.079**	
		[0.030]	[0.029]	
SD of real exchange rate	**0.021****	0.013		**0.014****
	[0.009]	[0.008]		[0.007]
Natural resource rents	**0.050*****	**0.048*****		0.013
	[0.004]	[0.016]		[0.019]
Bureaucratic quality			-0.181	
			[0.132]	
Fiscal rule		-0.264		
		[0.189]		
Number of observations/countries	1022/48	1022/48	1022/48	1022/48
Number of instruments	129	130	128	11
Specification tests (p-values)				
Arellano-Bond AR(1) test	0.008	0.010	0.014	0.001
Arellano-Bond AR(2) test	0.492	0.455	0.441	0.811

Notes:
1) The table reports one and two-step System GMM dynamic panel estimations carried out using the 'xtabond2' package in Stata (Roodman, 2009). Following Fatás and Mihov (2003), we measure volatility using the standard deviation of the residuals from estimated fiscal reaction function of the cyclically adjusted fiscal balance on its lagged value, lagged government debt, and lagged output gap.
2) All results in the table are presented in easily comparable format. Regression 1 (one-step GMM) have exactly the same specification as regression 4 (two-steps GMM).
3) Regressions include the following control variables - real GDP per capita, standard deviation of cpi, trade openness, domestic credit, standard deviation of stock market cap., government fragmentation, corruption, and crisis, which were found to be insignificant.
4) *, **, *** indicates significance at 10% ($P < 0.10$), 5% ($P < 0.05$) and 1% ($P < 0.01$) respectively.

Source: Authors' estimations.

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Euro-area labour markets: Different reaction to shocks?

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ABSTRACT

A small labour market model for the six largest euro-area countries (Germany, France, Italy, Spain, the Netherlands and Belgium) is estimated in a state space framework. The model entails, in the long run, four driving forces: trend labour force, trend labour productivity, long-run inflation rate and trend hours worked. The short run dynamics is governed by a VAR model including six shocks. The state-space framework is convenient for the decomposition of endogenous variables in trends and cycles, for shock decomposition, for incorporating external judgment, and for running conditional projections. The forecast performance of the model is rather satisfactory.

The model is used to carry out a policy experiment with the objective of investigating whether euro-area labour markets react differently to a reduction in labour costs. Results suggest that, following the 2008–2009 recession, moderate wage growth would significantly help delivering a more job-intense recovery.

JEL classification: C51; C53; E17; J21

Keywords: Labour market, Forecasting, Kalman Filter

1. INTRODUCTION

The main objective of this paper is to illustrate some key features of the euro-area labour markets, in particular how euro-area labour market differ when they are affected by the same type of shocks. This question is relevant both for monitoring labour market developments in the euro-area and for policy analysis, i.e. which reform fits better one country vis-à-vis another.

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For this purpose a small labour market model for the six largest euro-area countries – Germany, France, Italy, Spain, the Netherlands and Belgium – is estimated in a state space framework. The model consists of a set of labour market equations (labour force, labour demand, wage curve, production function, relative prices, hours worked) which are jointly estimated.

On the quantity side the model distinguishes between the intensive (hours worked per persons) and the extensive (persons employed) margins. On the price side it distinguishes between GDP and consumer price deflators. The model entails long-run dynamics and short-run fluctuations. The long-run dynamics is derived from strong theoretical restrictions which determine the pattern of trends (labour force, labour productivity, inflation rate and hours worked). The short-run fluctuations are driven by a homoscedastic vector autoregressive process.

In addition, the model presents a number of features which in our view makes it appealing in comparison with similar studies. Firstly, the same specification is estimated for all countries, which allows for a straightforward cross-country comparison of the different reactions to shocks. Secondly, the estimation technique used allows the joint estimation of long and short run dynamics, where both the trend and the cyclical component have theoretical underpinnings. Therefore, there is no need of detrending data by purely statistical methods prior the model estimation.[2] Thirdly, on the data front, the paper makes use of recently available hours worked series at a quarterly frequency compatible with national accounts.

While the model is primarily empirical, the long-run restrictions are consistent with a frictionless economy where a Cobb-Douglas production function is used to derive the desired level of employment by firms. On the labour-supply side, the long-run wage curve is consistent with a bargaining model where the real consumption wage depends on productivity and on the prevailing labour market conditions, which are captured by the rate of unemployment.

The model is cast in a state-space framework, which is very convenient for the decomposition of endogenous variables in trends and cycles, for the shock decomposition, for incorporating external judgments and for running conditional projections.

Focusing on developments during the past 10 years, this paper shows how the euro-area countries differ in terms of contribution of the long-term driving forces and the short-term shocks to key labour market developments.

Finally, the paper presents a policy experiment with the objective of investigating the different reactions of the euro-area labour markets to a reduction in labour costs. This question, which has been already addressed by a number of previous empirical papers, is still very relevant in light of the ongoing debate on the need for regaining cost competitiveness in the euro-area for two key reasons. Firstly, given current uncertain situation regarding the employment prospects after the 2009 recession, it is important to shed light on the employment implications of a change in labour cost. Secondly, the 2009 recession implied very different labour market reactions across countries in terms of employment and hours worked, partly due to nominal-wage inertia. Thus it is important to understand the implications of a labour moderation strategy on the two margins of adjustment. Given that the estimation horizon used in this paper is relatively up-to-date, i.e. from 1992 to 2009, and the model is able to distinguish between intensive and extensive margins of labour utilisation, the answers provided by the paper to the above policy questions represent a new feature.

The model does not contain features of typical long-term growth models (e.g. population growth) because its main purpose is to give a realistic representation of the current labour market developments rather than focusing on long-term trends.

[2] Canova and Ferroni (2009), who adopted a non-structural approach to detrending using a parsimonious econometric specification, recently confirmed by simulations that an incorrect specification of trends distorts the estimation of parameters of the cyclical part of macroeconomic models.

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The paper is organized as follows. Section 2 is devoted to a survey of related papers. Section 3 presents the model. The dataset and the estimation are discussed in Section 4, while the results of a policy experiment are present in Section 5. Section 6 concludes.

It should be noted at the outset upfront that, given the empirical nature of the model and the absence of 'deep' parameters in the sense of structural models, the word 'shock' is used through the paper in the spirit of the traditional macro-modelling environment.

2. THE LINK WITH RELATED LITERATURE

On the modelling side, the literature on small-scale labour market models is considerable. Small-scale labour market models or the supply-side block of large macro-models usually involve a production function equation, a labour supply relationship, a labour demand derived from the first order conditions of profit-maximizing or equivalently a cost-minimizing representative firm and a wage setting relation determined by a bargaining process between firms and labour unions (Layard et al., 1991). On the empirical side, one of the critical problems is the identification of labour demand and supply relationships.

Following Morgan and Mourougane (2005), the labour demand equation is identified by using the value added deflator at factor costs to compute real wages, which is the relevant deflator for firms. By contrast, the wage curve is identified by using the consumption deflator to compute real wages, which is instead the relevant deflator for households. Moreover, labour demand conditions, i.e. the unemployment rate, enter in the wage curve. Following Hansen and Warne (2001) our model entails also a short-term equation for relative prices, that is the consumer/producer price wedge. Such a wedge reflects, in particular, the impact of indirect taxes, terms of trade effects and relative bargaining powers. In the short run, the dynamics of producer prices and of nominal wages is based on a VAR form with identification restrictions.

A similar approach has been followed by Duarte and Marques (2009), where an empirical SVECM involving nominal wages, prices, the unemployment rate, productivity and import prices is estimated for the euro area and the US. The main finding of the paper is that wage dynamics are mainly determined by unemployment shocks in both economies but a significant role is also played by technology shocks in the US and by import price shocks in the euro area. This last result is particularly important as it suggests that in the euro area economies wages tend to be 'de facto' indexed to imported inflation.

On the data side, due to limited availability of reliable time series for hours worked, so far most of the existing studies on euro-area countries labour markets measure employment as the number of persons employed (see Mourre, 2006). In this respect, the fact that our model is able to distinguish empirically between the intensive margin (hours worked) and the extensive margin (jobs) represents an important innovation with respect to standard labour market models.

On the estimation side, our econometric approach follows very closely the analysis in Proietti and Musso (2007), where a multivariate structural time-series model is used for the estimation of potential output and output gap. As in Proietti and Musso (2007), the decomposition of time series to trend and cyclical component in this paper is model-based and hence does not depend on purely ad hoc statistical techniques. The difference between these two papers lies in the specification of the model used to separate trend and cycle.[3] Brůha (2011) presents a similar model for the Czech economy.

Finally, regarding the policy content of the paper, a large number of empirical and theoretical works have covered the issue of the relative gains/costs of wage moderation. In particular, partial equilibrium approaches indicate that labour cost moderation generally help employment

[3] Hjelm and Jönsson (2010) provide an overview of various approaches to filter the trend component from economic time series, including multivariate model-based approaches.

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creation (Pierluigi and Roma, 2008) and growth (Estevao, 2005), the same applies to simulations conducted with large macroeconomic models (Angelini et al, 2013), where the key mechanism at work is the competitiveness channel which leads to higher growth and employment. While this last study looks at the implication of nominal wage moderation, the two studies previously quoted refer to real-wage moderation. In our simulation the focus is on nominal wages, this is because for countries belonging to a monetary union the ability to achieve nominal-wage moderation is very important, especially in an environment characterized by moderate price developments.

3. MODEL

The dynamics of the model is basically composed of two parts: the long-run dynamics and short-run fluctuations. The long-run dynamics is derived from strong theoretical restrictions and it provides a discipline on trends in modelling variables. The short-run dynamics then enriches the structure of the model and makes it possible to use the model for forecasting and shock decomposition exercises.

In general terms, any model variable x_t is given as a sum of the trend component and the cyclical component[4]

$$x_t = \bar{x}_t + \tilde{x}_t$$

where $\bar{x}_t$ is the trend component, and $\tilde{x}_t$ is the cyclical component.

3.1. The long-run dynamics

The equations describing the long-run dynamics are given as follows, where all variables are in logs:

$$\bar{y}_t = \bar{e}_t + \bar{h}_t + \varpi_t^z \tag{1}$$

$$\bar{e}_t + \bar{h}_t = \bar{y}_t - (\bar{w}_t - \bar{p}_t) \tag{2}$$

$$\bar{w}_t - \bar{q}_t = -\gamma_1\left(\bar{l}_t - \bar{e}_t\right) + \gamma_2\left(\bar{y}_t - \bar{e}_t - \bar{h}_t\right) + \alpha_2 \bar{m}_t \tag{3}$$

$$\bar{p}_t = \varpi_t^p \tag{4}$$

$$\bar{l}_t = \varpi_t^l + \alpha_1 \bar{m}_1 \tag{5}$$

$$\bar{h}_t = \varpi_t^h \tag{6}$$

In Eq. (1), (2) and (3) $\bar{y}_t$ is the trend output, $\bar{e}_t$ is the trend number of persons employed. In Eq. (1), (2), (3) and (6) $\bar{h}_t$ is the trend number of hours worked per employee. In Eq. (2) and (3) $\bar{w}_t$ is the trend in the nominal compensation per total hours worked. In Eq. (2) and (4) $\bar{p}_t$ is the trend GDP deflator at factor costs; in Eq. (3) $\bar{q}_t$ is the private consumption deflator (which is assumed to follow the same trend as $\bar{p}_t$), $\bar{l}_t$ is the labour force, which is the sum of employed and unemployed, and $\bar{m}_t$ is the trend in the net immigration flows, which affects the labour force. In Eq. (1), (4), (5) and (6) ϖ_t^k with $k = z, p, l, h$ denote trends in productivity, price level, labour force, and hours worked.

[4] All variables in this paper are considered in logs unless otherwise stated. All variables are also seasonally adjusted.

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The long run specification of the model is stylised. In particular, Eq. (1) suggests that the production function is formulated as a relation describing average labour productivity which follows, in the long run, the productivity trend ϖ_t^z. Eq. (2) is derived from the first order conditions of the Cobb-Douglas production function and expresses the desired total amount of hours worked $\left(\bar{e}_t + \bar{h}_t\right)$ as a function of the level of output and the real product wage $(\bar{w}_t + \bar{p}_t)$. The consumer real wage $((\bar{w}_t + \bar{q}_t))$ is determined by a bargaining process between firms and labour unions. The outcome of this process is described as a relationship between the consumer real wage, average real productivity and unemployment in Eq. (3). It is assumed that net immigration is able to affect the consumer real wage by reducing the bargaining power of unions.

The trend price level (ϖ_t^p) is assumed to be the same for the GDP and consumption deflators (Eq. (4)). Eq. (5) says that trend domestic labour force ϖ_t^l is affected by immigration. The long-run dynamics of hours worked is determined by its trend ϖ_t^h, which reflects slowly-moving institutional features of the economy; see Eq. (6). The structural form implies the following long run elasticities:

	ϖ_t^l	ϖ_t^z	ϖ_t^p	ϖ_t^h
$\bar{l}_t$	1	0	0	0
$\bar{y}_t$	1	$1 + \dfrac{1-\gamma_2}{\gamma_1}$	0	1
$\bar{e}_t$	1	$\dfrac{1-\gamma_2}{\gamma_1}$	0	0
$\bar{w}_t$	0	1	1	0
$\bar{p}_t$	0	0	1	0
$\bar{q}_t$	0	0	1	0
$\bar{h}_t$	0	0	0	1

(7)

Given that the parameters γ_1 and γ_2 are not separately identified, only their non-linear combination $\dfrac{1-\gamma_2}{\gamma_1}$ could be estimated.

To model trends ϖ_t^k we modify the approach by Harvey and Jaegger (1993). Harvey and Jaegger (1993) propose an I(2) process for filtering trends from economic time series, which is defined as follows:

$$\theta_{1t} = \theta_{1t-1} + \theta_{2t-1} + \sigma_1 \eta_{1t}$$

$$\theta_{2t} = \theta_{2t-1} + \sigma_2 \eta_{2t}$$

where η_{it}, $i = 1, 2$ are independent i.i.d. white-noise processes. Standard errors σ_1, σ_2 determine the smoothness of the filtered trend. In particular, if σ_1 is small, the trend θ_{1t} is relatively smooth.

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It is interesting to note that the Leser (1961) filter (later 'rediscovered' in economics as the HP filter) is optimal for processes $x_t = \theta_{1t} + v_t$, provided that $\sigma_1 = 0$ and v_t is an i.i.d. white noise sequence. Although, the Harvey-Jaegger model can be used to filter out smooth trends, it is not suitable for forecasting.[5]

We therefore opt for an I(1) process to model ϖ_t^k, since this would imply that the long-run growth rate in the variables is a stationary process. This is a rather plausible feature especially for productivity and the price level.[6] In more details, we assume the following ARIMA (3,1,0) process, which can produce smoothed growth trends slowly varying around a long-run value:

$$\varpi_t^h - \varpi_{t-1}^h = \theta_{1t}^k, \tag{8}$$

where θ_{1t}^k follows a stationary process:

$$\theta_{1t}^k = \rho_1 \theta_{1t-1}^k + \theta_{2t-1}^k$$

$$\left(\theta_{2t}^k - \mu^k\right) = \rho_2 \left(\theta_{2t-1}^k - \mu^k\right) + \theta_{3t-1}^k$$

$$\theta_{3t}^k = \rho_3 \theta_{3t-1}^k + \varepsilon_t^k$$

with ε_t^k being an i.i.d. white noise sequence, and $0 \leq \rho_i < 1$ for $i \in \{1,2,3\}$. The fact that stochastic innovations enter directly only into the third equation implies that θ_{1t}^k follows a slowly varying smooth process.

The interpretation of model (8) is the following: θ_{1t}^k can be considered as the trend growth rate of the variable ϖ_t^k, which moves around a target θ_{3t}^k. This target grows in a steady state by μ^k and is shocked by an invertible MA(1) process, represented by θ_{3t}^k. Note that the process is similar to the one used in Carabenciov et al. (2008) (and in related IMF-based models) to filter low-frequency movements in output and unemployment. The difference with respect to Carabenciov et al. (2008) is that we add a process θ_{3t}^k, which adds an additional flexibility to the spectral properties of the low-frequency component.

3.2. Formulation of the state-space model

The short run dynamics of the model is governed by a VAR(1) process. Both short and long term components are then combined in a single state-space model, which is then used for model estimation and simulation. The observation variables are transformed to annualised quarterly growth rates.[7]

Hence the observed growth rate of the labour force is given by:

$$l_t - l_{t-1} = \left(\bar{l}_t - \bar{l}_{t-1}\right) + \left(\hat{l}_t - \hat{l}_{t-1}\right) + \theta_{1t-1}^l + \left(\hat{l}_t - \hat{l}_{t-1}\right)$$

[5] The problem may occur when the filter identifies a sign change in θ_{2t}. Such a change is then permanent on forecast. As an example, this may happen for trend productivity during a huge recession. If a model identifies in the last period that $\theta_{2t} < 0$, it will then predict an indefinite decline in productivity, which would mean that the model would not be able to forecast any growth recovery no matter how far in the future. Unless one believes in such a doomsday scenario, this is clearly an implausible feature for forecasting.

[6] At least in the case of a well anchored monetary policy.

[7] Here, we assume that we measure and filter only growth rates. In models where additional restrictions (in the form of e.g. accounting identities or equilibrium conditions) were imposed, it would be beneficial to measure also the level of variables directly. However, we do not have additional restrictions here and thus no efficiency is lost by measuring the growth rates only.

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The observed growth rate of output is given by:

$$y_t - y_{t-1} = (\bar{y}_t - \bar{y}_{t-1}) + (\hat{y}_t - \hat{y}_{t-1}) + \theta^l_{1t} + \left(1 + \frac{1-\gamma_2}{\gamma_1}\right)\theta^z_{1t} + (\hat{y}_t - \hat{y}_{t-1})$$

where the second equality follows by virtue of long-run elasticities[8]. Similarly, the observed employment is given by:

$$e_t - e_{t-1} = (\bar{e}_t - \bar{e}_{t-1}) + (\hat{e}_t - \hat{e}_{t-1}) + \theta^l_{1t} + \left(1 + \frac{1-\gamma_2}{\gamma_1}\right)\theta^z_{1t} + (\hat{e}_t - \hat{e}_{t-1})$$

The observed nominal wage inflation $\pi^w_t = w_t - w_{t-1}$ is given by:

$$\pi^w_t = \bar{w}_t - \bar{w}_{t-1} + \hat{\pi}^w_t = \theta^p_{1t} + \theta^z_{1t} + \hat{\pi}^w_t$$

Analogous formulae apply for changes in the GDP and consumption deflators:

$$\pi^y_t = \theta^p_{1t} + \hat{\pi}^y_t$$

$$\pi^c_t = \theta^p_{1t} + \hat{\pi}^c_t$$

Measurement noise has been not included in the system.

4. DATA AND ESTIMATION OF THE MODEL

As compared to previous empirical work on labour market models, the dataset used is very up-to-date. Data have quarterly frequency and span from 1992Q1 to 2009Q4 for Belgium, Germany, Spain, France, Italy and the Netherlands. The choice of this sample period has been dictated both on the basis of statistical and economic grounds. On the statistical side, this sample period excludes the data problem related to German unification and to a number of missing back data for some countries, in particular hours worked. On the economics side, this period excludes the strong cost and price disinflationary process undergone during the 1980s by most euro-area countries. While this implies reducing the volatility of the series and thus limiting their explanatory power, the choice of limiting the estimation to this relatively recent sample period is more economically founded, as it excludes the disinflationary period of the period of the previous decade. All data used are adjusted for seasonality.

Unit wages are measured as gross compensation (including social security contributions) per hours worked and labour productivity is measured as real GDP per total hours worked. Given that the unemployment rate features also the model, which is typically measured as number of individuals, a distinction is made between the intensive and extensive margins. Real GDP and its deflator are measured at factor costs. The private consumption deflator is used as a measure for consumer prices. The wedge between the consumption and GDP deflators captures the effect of taxes and administered prices, as well as those of relative import prices on wage bargaining. In this regards external shocks are feeding throughout the model via this wedge. As typically handled in the empirical literature the identification problem of the labour demand and supply equations is

[8] Results in Table 7 imply that $\bar{y}_t = \varpi^l_t + \left(1 + \frac{1-\gamma_2}{\gamma_1}\right)\varpi^k_t$.

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resolved by assuming that real product wages is what matters for firms (labour demand) and real consumer wage is what matters for employees or unions (wage curve).

Regarding some properties of the data, it is worth mentioning that quarterly series of hours worked have been only very recently made available for the euro-area countries by the national statistical offices. These data, while new, confirm the long term downward trend in the estimated data available for the annual frequency. As is evident from Figure 1, the employment gains across the euro-area countries were mainly determined by the creation of new jobs, while the working time per person has declined substantially during the past 20 years. As reported in Leiner-Killinger et al. (2005), the decline in hours worked is attributable to the increased use of part-time working arrangements, which is often related to the greater number of women entering the labour market, to institutional factors such as tax wedges which create disincentive to work, or to specific policy measures including changes in working time regulations, such as the introduction of the 35-hour week in France in 2000. During the 2009 recession, in some euro-area countries the reduction in total hours was a temporary phenomenon, primarily driven by the introduction of government-sponsored short-time work measures and flexible working time arrangements (such as working-time accounts).

Labour market reforms pursued in the early and mid-2000s in the largest euro-area countries aimed at increasing the flexibility for new hires to enter and leave new occupations have provided strong incentives for part-time jobs at the expense of the total number of hours worked. Given this decoupling between hours worked per person and employment it appears very relevant also for policy purposes to measure the unit wage as compensation per total hours worked rather than as per person employed, which is instead the typical approach in earlier empirical work on labour demand and supply (see for example Mourre, 2006). In particular, during the 2009 recession the dynamics of hourly compensation remained rather strong, while that of compensation per employee has moderated significantly.

The model described in Section 3 can be rewritten according to the following state- space form:

$$x_t = Ax_{t-1} + M + \Sigma_x \epsilon_t \tag{9}$$

$$y_t = Cx_t \tag{10}$$

where the state vector contains the vectors of the long-run growth rates θ^k_{1t}, their drivers θ^k_{2t}, θ^k_{3t}, cyclical parts of the model; the matrix A contains the long-run block based on the process (8) and the VAR block, which drives the short-run dynamics; the matrix M contains μ^x_t in the appropriate locations; $\Sigma_x \epsilon_t$ maps structural innovations to θ^k_{3t}, and to the cyclical part; and finally the matrix C adds the two components to the vector of the observable variables y_t. The part of this matrix, which corresponds to the long-run dynamics, is based on the long-run elasticities reported in (7).

The Kalman filter is used for state filtering and smoothing, forecasting, likelihood evaluation and shock decomposition. The formulae for filtering, smoothing, and likelihood evaluation are fairly standard (e.g. Harvey, 1989). Given the smoothed estimate of the state $x_{T|T}$ and its covariance matrix $P^x_{T|T}$ (T denotes the last observation), the h-step (unconditional) prediction of $y_{T+h|T}$ can be computed as:

$$y_{T+h|T} = C\left[A^h x_{T|T} + \left(\sum_{g=0}^{h} A^g\right)M\right]$$

and its covariance matrix can be derived as follows:

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$$P^{y}_{T+h|T} = CP^{x}_{T+h|T}C'$$

where the state prediction covariance matrices $P^{x}_{T+h|T}$ satisfy the recursion:

$$P^{x}_{T+h+1|T} = AP^{x}_{T+h|T'}A' + \Sigma_x \Sigma_x'$$

with the initial condition given by the Kalman filter output $P^{x}_{T|T}$.

The conditional forecast can be also easily derived. To condition the forecast on a set of variables, it is sufficient to run the filter on the model with a suitably redefined observation matrix C. The approach used for implementing conditional forecast maintains trends fixed on the unconditional projection. In this way, we can attribute the difference between conditional and unconditional forecast to the cyclical part of the model. Also, following Koopman and Harvey (2003), the Kalman filter can be 'inverted' to inquire how observations in each series translate to the model assessment of trends and cycles. Finally, the model is used to perform shock decomposition, which is computed as follows: based on the smoothed states $x_{t|T}$, we can recover the smoothed residuals $\epsilon_{t|T}$. The shock contribution to the ith observable variable from the jth shock is defined as $\Sigma_{s\geq=0} i^{ij}_{s} \epsilon^{j}_{t-s|T}$ where $\epsilon^{j}_{t-s|T}$ is the jth element of $\epsilon_{t-s|T}$ and i^{ij}_{s} is the (i, j) element of the impulse response $i_s = CA^s\Sigma_x$. Note that our definition of the shock decomposition cumulates the effects of current and past shocks (the alternative decomposition could be defined in terms of current shocks plus the effect of the initial conditions).

We estimate the model using a pseudo-Bayesian approach (as in Hong and Chernozhukov, 2003): we maximize the likelihood with some prior imposed on long-run growth rates[9] and on selected signs of impulse responses for the short-run dynamics. It is interesting to mention, that only for Italy does the prior affect the estimation results. For other countries, the mode of the posterior distribution would be almost equal to the maximum likelihood estimator.[10] The prior on standard errors of innovations are flat (uninformative) and hence the same for all countries.

Figures from 2 to 7 show recursive point forecasts for the six euro-area countries since 2004. All the charts show annualized quarterly changes. Recursive point forecasts show that in all cases the model can track the short-term dynamics rather well, with the exception of the labour force in Germany, France, Italy, and the Netherlands. This may be due to the fact that in these countries the labour force has shown very little procyclicality in the recent past, due to structural factors, such as population aging. Figures 9 to 14 compare the relative accuracy of the model forecasts (denoted as the BPS model, based on the first letters of authors' names) with forecasts generated by the 'random walk' model, by unrestricted VAR(1) and unrestricted VAR(2) processes[11] at forecast horizons from 1 to 8 quarters. The figures display the root mean square errors (RMSE) relative to BPS; the RMSE of the model presented being normalized to one. Typically, for most countries, BPS does a better job for nominal wages and inflations, while for the labour force the unrestricted VARs seem to be better. The ordering for the rest of variables is inconclusive.

Although in the case of some variables the forecast performance of our model is as good as VARs models, the advantage of the state-space framework used should be stressed. The state-space formulation allows one easily to make conditional forecasts or to incorporate external

[9] All simulations reported in this paper are taken with respect to the mode of the posterior distribution.

[10] For Italy, the maximum likelihood estimator would result in extremely low standard errors of innovations in the process (8). This would mean that the long-run growth rates would be almost constant, which is a feature we do not like. Such a feature is especially counterfactual when the model is used for forecasting during a crisis for it would predict a rapid recovery.

[11] A reader may ask why we have chosen the VAR order of the maximum lag 2. The reason is that VARs with higher lags are better at the very short forecast horizons (1 quarter), but their forecasting ability deteriorates rapidly for longer forecast horizons (for horizons greater than 3, they are much worse than any of the four models considered). This effect is likely due to over-parametrization of higher-order VARs. The reader may want to consider the lucid discussion by Tiao and Xu (1993) for intuition why some models can be 'good' at short forecast horizons, but fail completely at longer horizons.

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judgements. For example, if the forecaster has an extra piece of information (say from sectoral experts) about the likely evolution of only some of the model variables, then they can run conditional forecasts based on this piece of information. It is sufficient to redefine the matrix C from Eq. (10), as it just means to delete rows of the matrix C corresponding to variables for which the information is not available. The same approach can be used if some variables are measured with a lag or lead comparing to the other variables.

Similarly, if the forecaster doubts regarding the real time release of some observations (for example, the national-account wages are too different from the census-based wages) and fears that the figure may be subject to significant revisions, they could add a measurement noise to Eq. (10) and run the filter and forecast with the last observation subject to the measurement noise.

Furthermore, the state-space framework can incorporate the 'expert-information' about unobserved trends. For example, if the model yields the implausible decomposition of a series to trend and cyclical components, it is possible to force the model to a more plausible decomposition by a suitable expansion of the observation matrix C for a given period. By this, the expert can impose their preferred decomposition for one or more series.

Finally, there is a notorious issue of the end-point bias, i.e. excessive trend revisions due to new data. Our model applied to the data exhibits lower revisions comparing to the mechanical application of the HP filter. In the literature, this is often explained as the virtue of the state-space formulation and the Kalman filtering[12], but this is incorrect. As shown by Andrle (2013), the 'end-point bias' of a filter (i.e., the excessiveness of trend revisions) depends on how well the assumed model can forecast future observations rather than on the filtering technique (Kalman filter versus least squares). It turns out that the trend formulation chosen in this paper can produce more reasonable forecasts than the I(2) Harvey-Jaegger process implicitly assumed by the HP filter (see Footnote 5 above), and this is the reason for alleviating excessive trend revisions.

Figures 15 to 20 show the historical shock decomposition for four variables. The left column displays the decomposition to the model trends and to the cyclical component, while the right column shows the decomposition of the cyclical part of the variable. Also in this case the charts show annualized quarterly changes.

Focusing on the left column of the figures 15–20, from a cross-country perspective it is interesting to note that the model reading of the 2009 crisis is that of an almost entirely cyclical episode, i.e. trend developments have not been affected by the crisis. The only exception is Spain, were a clear downward movement in employment and an upward movement in hours worked can be observed. As regards trend wages the countries can be divided in two groups: in the first group – Germany, France and the Netherlands – trend wages are equally explained by trend inflation and productivity; in the second group – Italy, Spain and Belgium trend wages are predominantly explained by trend inflation.

As to the evolution of the cyclical component during the past 10 years, Germany is the only country which appears characterized by an opposite movement of employment and labour participation shocks in explaining the dynamics of employment and hours worked. Indeed, across the euro-area countries, Germany is the country which witnessed the smallest employment creation between 1999 and 2008 and a sharp downward trend in the labour force. In other words the chart suggests that in Germany, employment creation was demand- and not supply-driven. As to the shock decomposition of the cyclical component of German GDP, this has been mainly dominated by a productivity shock. Indeed, a key issue for the German economy is that many years of productivity gains were translated only very marginally and very late into employment creation. As to the shock decomposition of the cyclical component of nominal wages in Germany, this has been mainly driven by participation and wage shocks.

[12] E.g. Proietti and Musso (2007) claim that the state-space model can alleviate the bias *because of the adaptation property of the Kalman smoother at the end of the sample.* Our working paper version also contains a similarly incorrect statement.

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As regards, the evolution of the cyclical components in France, all shocks appear to have contributed to the downward adjustment of employment during the recession. In Italy, according to the model, relatively favourable internal terms of trade could explain the benign employment movements in the early 2000s and could offset a larger fall of employment in the 2009 crisis, which was driven by productivity and employment shocks.

Finally in the case of Spain, the fall in employment between 2008 and 2009 is explained by a decline of trend labour force. In the case of trend output growth such a decline is compensated for by an increase in trend productivity growth. This result, which of course entails a high degree of uncertainty, would suggest that a rebalancing of the supply-side determinants of growth is taking place in Spain. Such a rebalancing might deliver, if persistent, a more sustainable growth model for the Spanish economy.

5. A POLICY EXPERIMENT: THE EFFECT OF WAGE MODERATION ON THE JOB-RICHNESS OF THE ECONOMIC RECOVERY

The 2009 recession, which is covered in the dataset used for the estimation, has led to very different employment responses across the euro-area countries. In particular, Germany and to a lesser extent Italy, Belgium and the Netherlands have witnessed a significant degree of labour hoarding, stronger than in previous recessions, while Spain saw exceptionally strong labour shedding. While labour hoarding is a common characteristic across the euro-area countries, the particularly strong resilience of the labour market to the sharp economic downturn was mainly due, in the first group of countries, to the extensive use of special measures to support employment. By contrast, the employment losses observed chiefly in Spain were related to the sharp downward correction of a strongly labour-intensive sector, i.e. construction, in an environment characterized by very loose firing conditions (due to a high rate of temporary contracts). In light of this heterogeneity in labour market adjustments across euro-area countries, the employment prospects in a recovery scenario appear highly uncertain, as it might be the case that firms would downwardly adjust employment once special schemes to keep jobs are expired or that they would gradually return to higher levels of hours worked per person, waiting for long before new job opportunities are created. On the other hand, it might also be the case that in those countries witnessing strong labour shedding, the recovery may provide a relatively stronger impulse to employment.

This section tries to answer the following policy question: to what extent could a stronger degree of hourly wage moderation than that recently witnessed strengthen the job-richness of the economic recovery after the 2009 recession?

The simulation exercise consists of quantifying the different elasticities of a 1% drop in the unit wage level across euro-area countries. These elasticities are derived by taking the difference between the unconditional forecasts delivered by the our model in a two-year horizon and the conditional forecasts, where a 1% drop in the wage level in the course of the first year (2010) has been assumed. Such a drop is obtained by reducing the wage rate in each quarter of 2010 by a proportional amount leading to a 1% fall with respect to the baseline level for the year as a whole. The drop in the level of wages is permanent, i.e. no unwinding has been implemented in the subsequent year. The simulation results are reported in Figure 8.

In general, this empirical model confirms the gains in terms of higher employment which could be achieved via wage moderation even in the short-run. However, the reaction of such a wage shock entails different implications in terms of margins of adjustment. It appears, in particular, that gains both in hours worked and persons employed could take place in such a scenario in Germany, France, the Netherlands and, after two-years, in Belgium. By contrast the same shock would induce, as in the case of Spain, very strong employment creation, especially in the second

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year after the shock, at the expense of hours per person, leading in any case to an overall gain in terms of total hours worked.

In the case of Spain, the finding of a trade-off between the evolutions of the two margins of adjustment can be explained by looking at past behaviour of the two variables. The Spanish unemployment rate has shot up by almost 10 percentage points since the beginning of the 2009 recession, while compensation has remained almost unaffected by the sharp change in labour market conditions. This feature has been found also in other empirical cross-country works (see Pierluigi and Roma, 2008). Such a huge employment correction has been partially driven by the burst of the housing bubble and partially by the strong bias towards fixed-term contracts. The model reading of such a situation is that even a small decline in the wage rate would strongly impact on employment and would also lead to a strong fall in the labour force. An important caveat is that this rather stylised model cannot capture sectorial adjustments and therefore tends to over-weigh the possibility for the unemployment rate to swiftly return to pre-crisis levels.

In the case of Italy the wage shock leads to a rather small reaction of hours worked and a small and negative reaction of employment in persons. The sum of the two margins imply an overall almost nil impact of the negative wage shock. The result is related to the very weak empirical link between wage dynamics and hour developments in Italy, as it emerges from the very small reaction of the labour demand to labour costs developments (see Pierluigi and Roma, 2008 and European Commission, 2006).

All in all, looking at the aggregate variable (EA6), one can conclude from this exercise that wage moderation would certainly help employment creation and – to a lesser extent – an upward adjustment in hours. In the case of Spain, the results suggest that a moderation in the dynamics of the wage rate would be highly beneficial for preventing further employment losses. Results are in line with other studies, based on larger scale models, where a negative wage mark-up shock is used to replicate an increase of labour market flexibility (see Gomes et al. 2011 and Angelini et al., 2013).

6. CONCLUSIONS

This paper presents a new macro tool for monitoring and forecasting labour market developments across the six largest euro-area countries. The model is primarily empirical but relies on theoretical underpinning in the derivation of the trends. The forecasting properties of the estimated model are satisfactory as they generally improve on first and second order VAR models and random walk processes.

The paper also shows that labour market adjustments differ substantially across euro-area countries, as it emerges from the contributions of the long-term drivers and short-term shocks to key labour market developments.

Finally, the model is used to assess the employment impact of reduction in the nominal hourly wage rate. The results of this policy experiment would suggest that in an environment characterized by significant labour hoarding, achieving moderate wage growth significantly helps delivering a more job-intense recovery. The simulations also show that countries tend to differ in their adjustment of hours worked versus job creation in response to a nominal wage cut, most likely in relation to the different institutional settings. In Spain, wage moderation appears particularly beneficial for preventing further employment losses after the sharp labour shedding witnessed in 2009. By contrast, in France, Italy and Belgium a lower nominal wage rate triggers a higher response of hours worked rather than jobs. In Germany and the Netherlands both margins of adjustment are equally positively affected by a nominal wage reduction.

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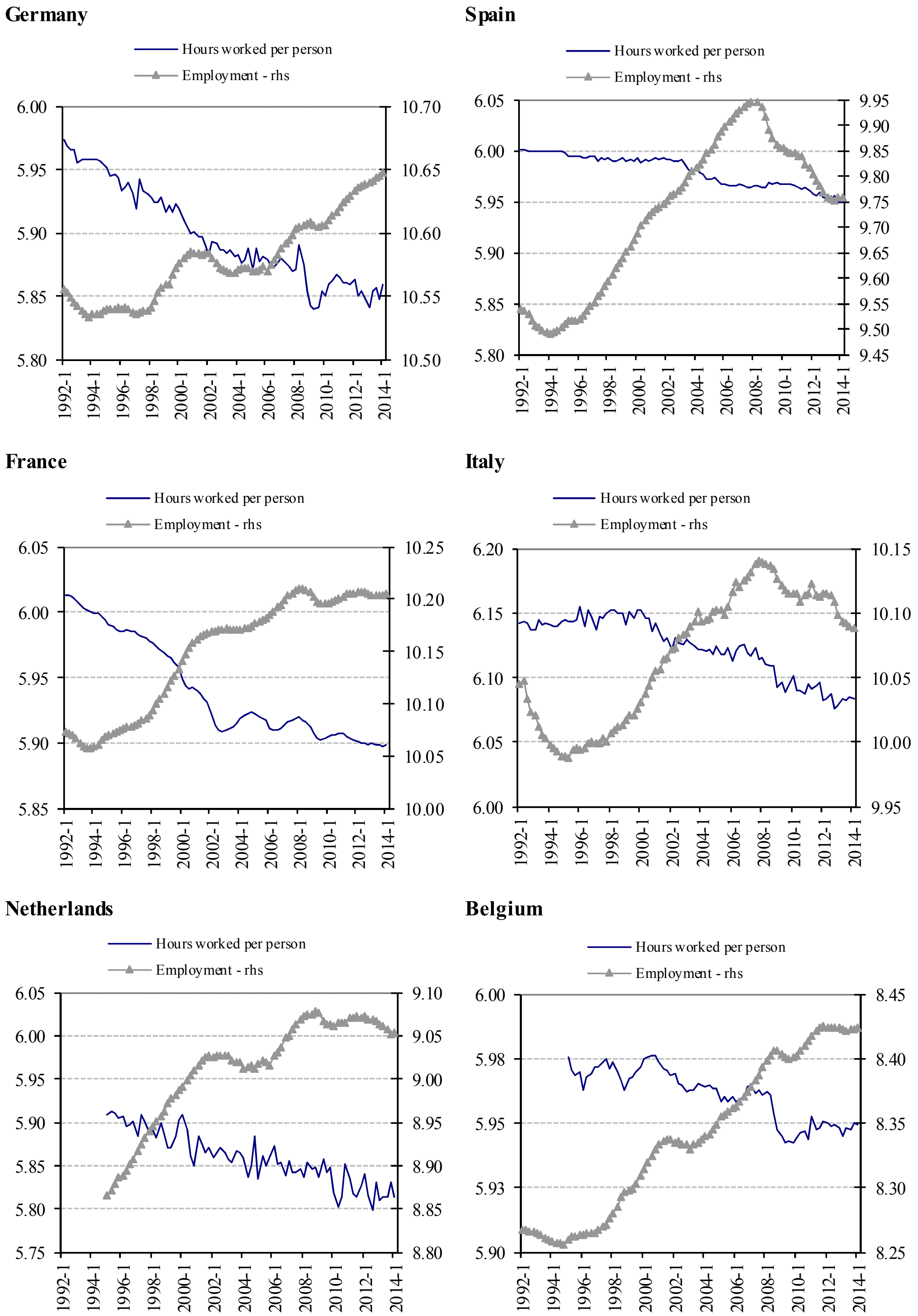
APPENDIX

Figure 1
Hours worked and employment developments across euro-area countries



Note: data shown in logs. In the case of NL and BE data for hours worked start in 1995Q1.

Source: Authors' calculation on Eurostat data.

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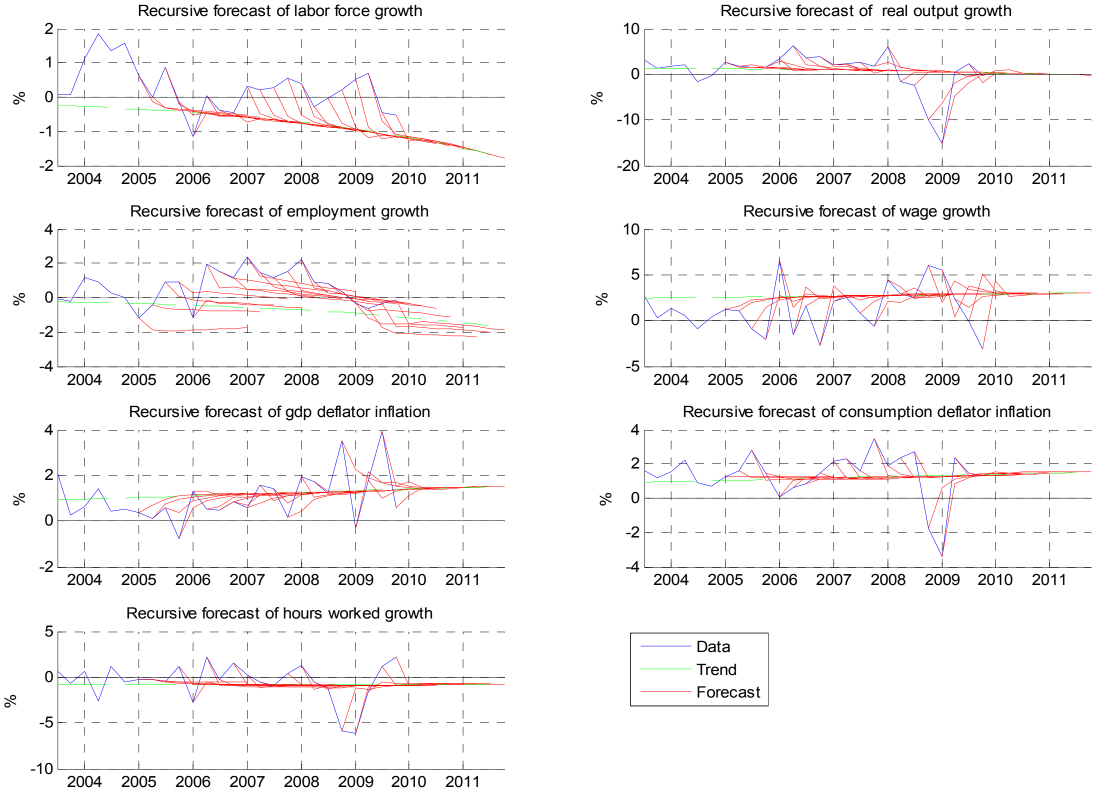
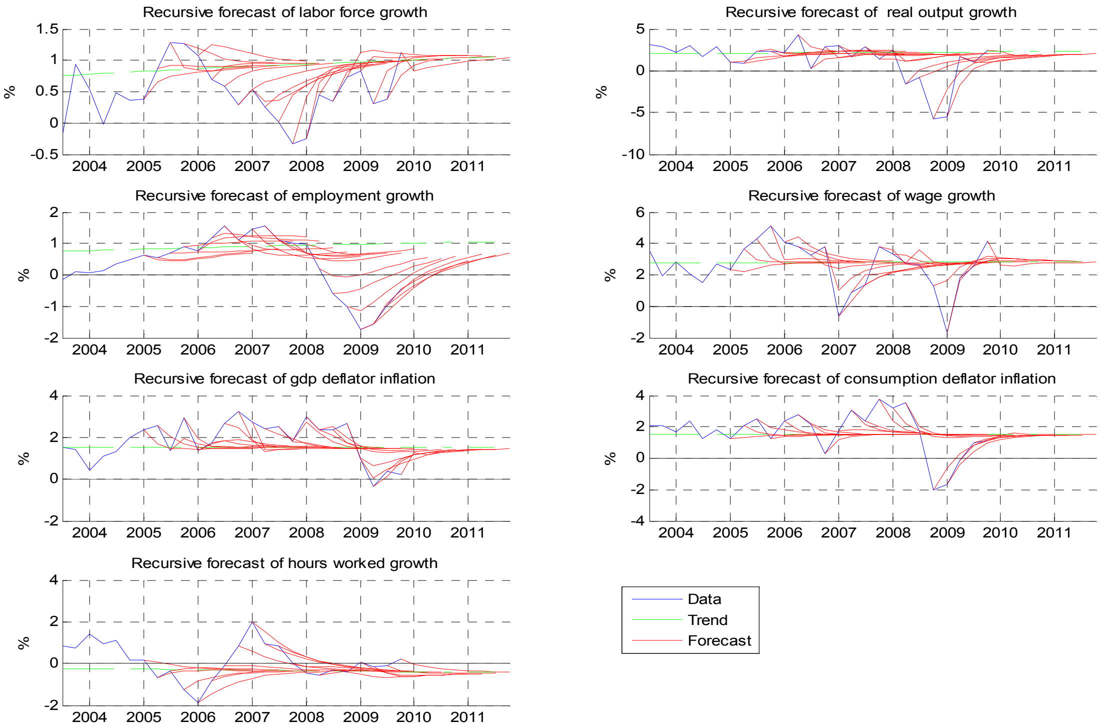
Figure 2
Germany – recursive forecast



Source: Authors' calculation.

Figure 3
France – recursive forecast



Source: Authors' calculation.

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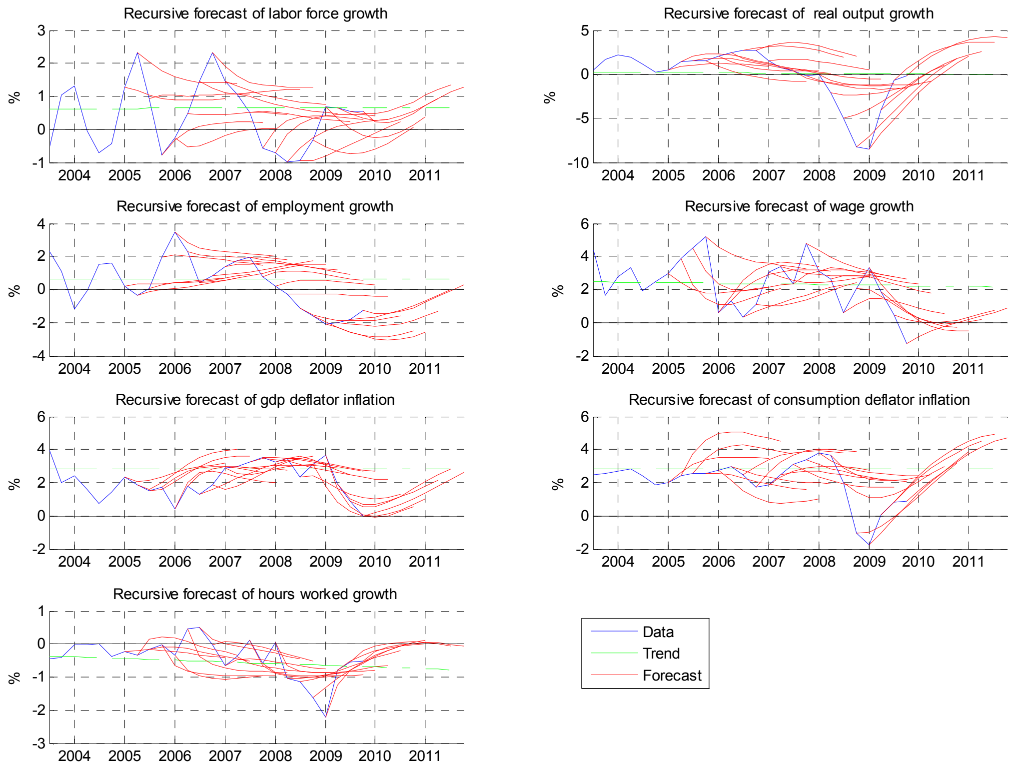
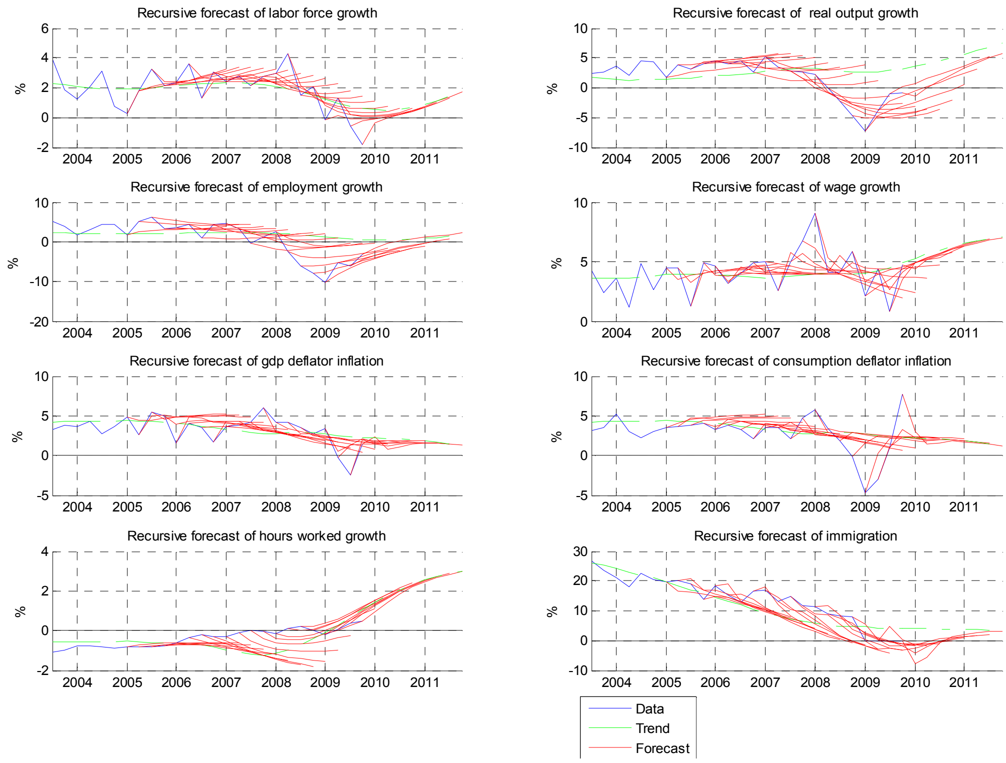
Figure 4
Italy – recursive forecast



Source: Authors' calculation.

Figure 5
Spain – recursive forecast



Source: Authors' calculation.

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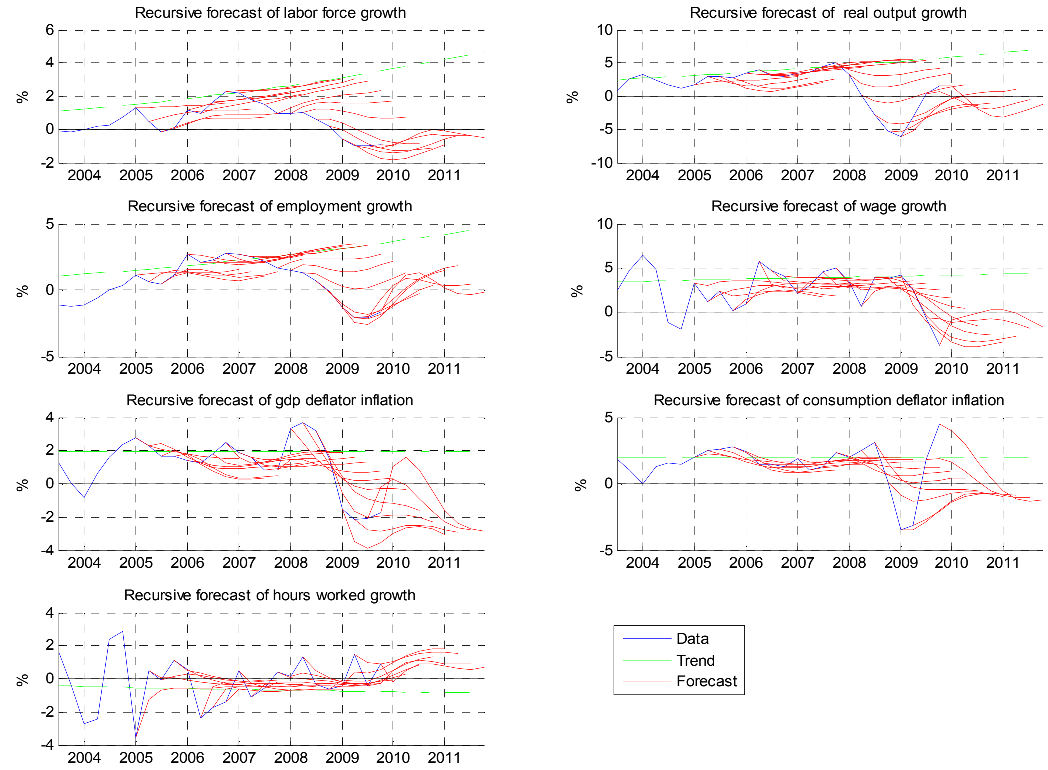
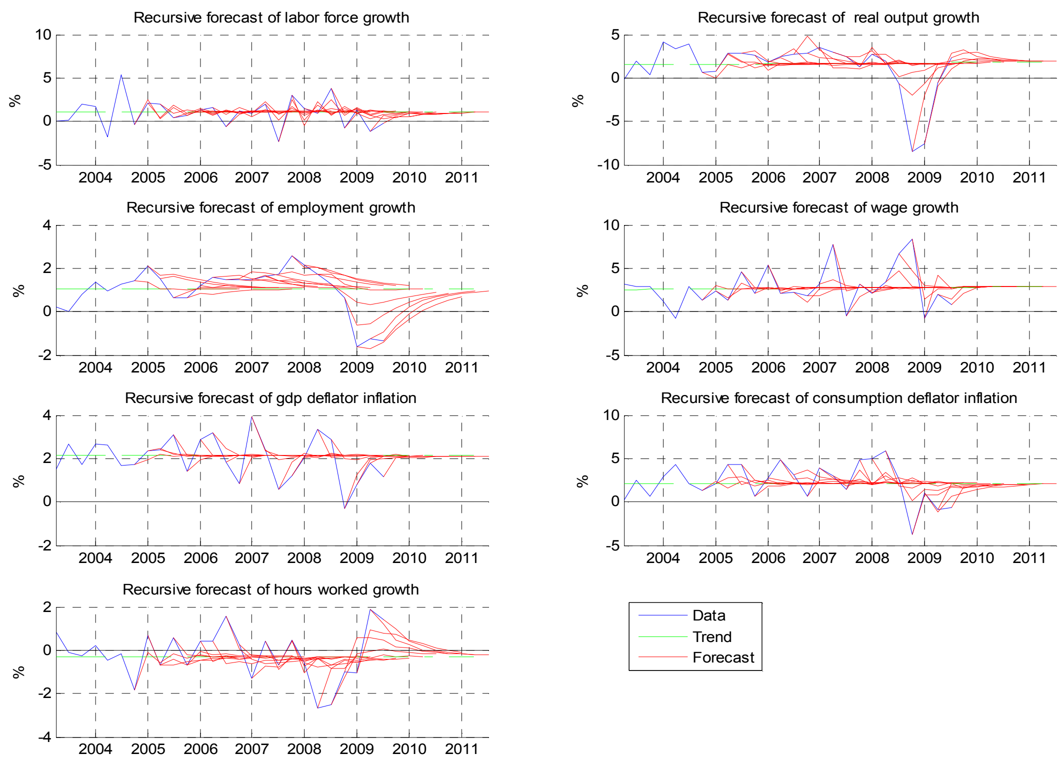
Figure 6
The Netherlands – recursive forecast



Source: Authors' calculation.

Figure 7
Belgium – recursive forecast



Source: Authors' calculation.

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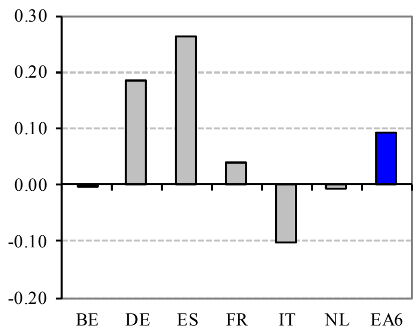
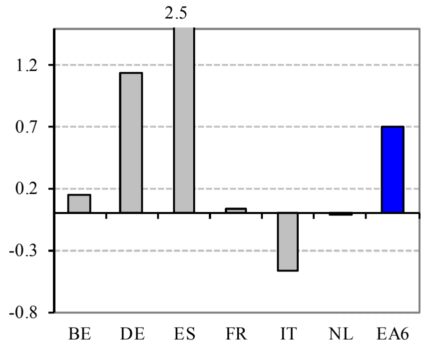
Figure 8
Impact of a 1% drop in hourly compensation

Employment in persons

(a) 2010

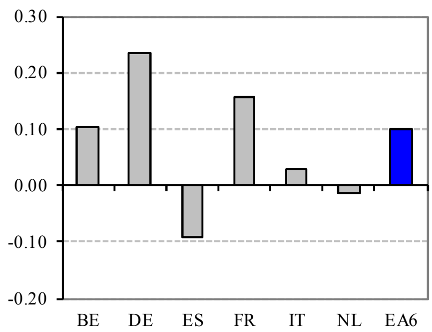


(b) 2011 (cumulated)

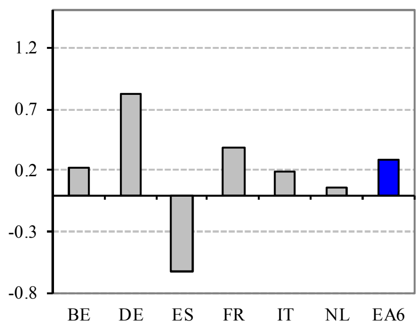


Hours worked per person

(a) 2010

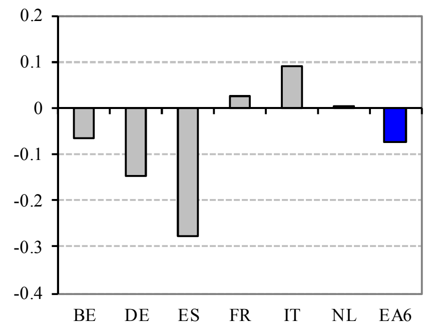


(b) 2011 (cumulated)

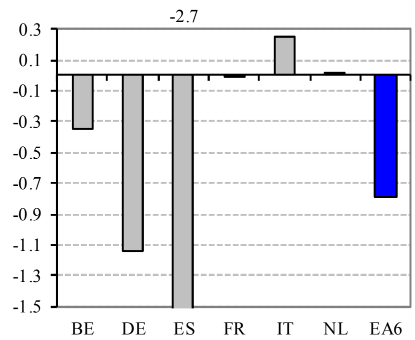


Unemployment rate

(a) 2010



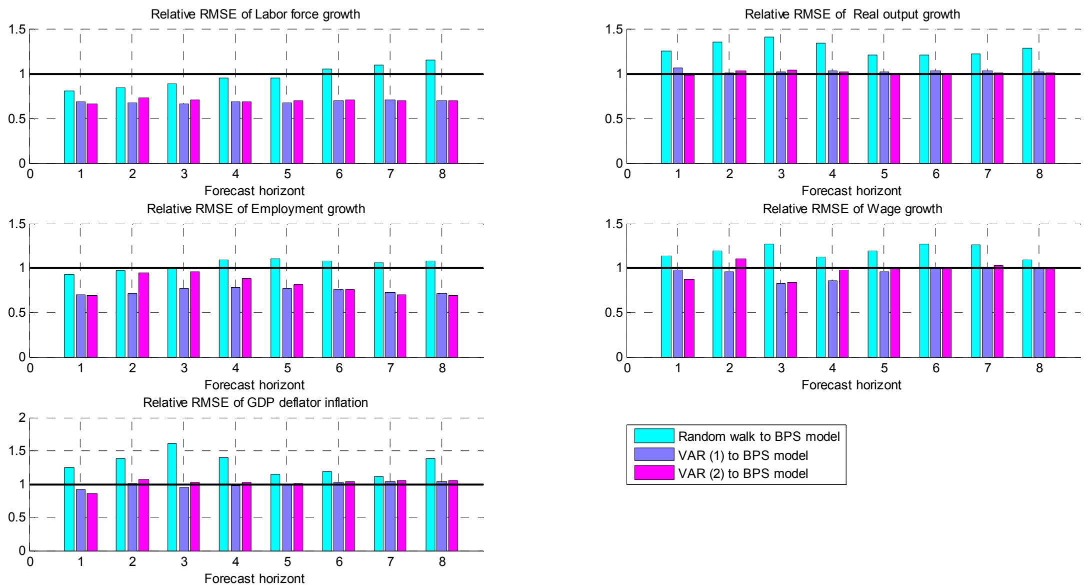
(b) 2011 (cumulated)



Note: EA6 is the weighted average of the six euro-area countries.
Source: Authors' calculation.

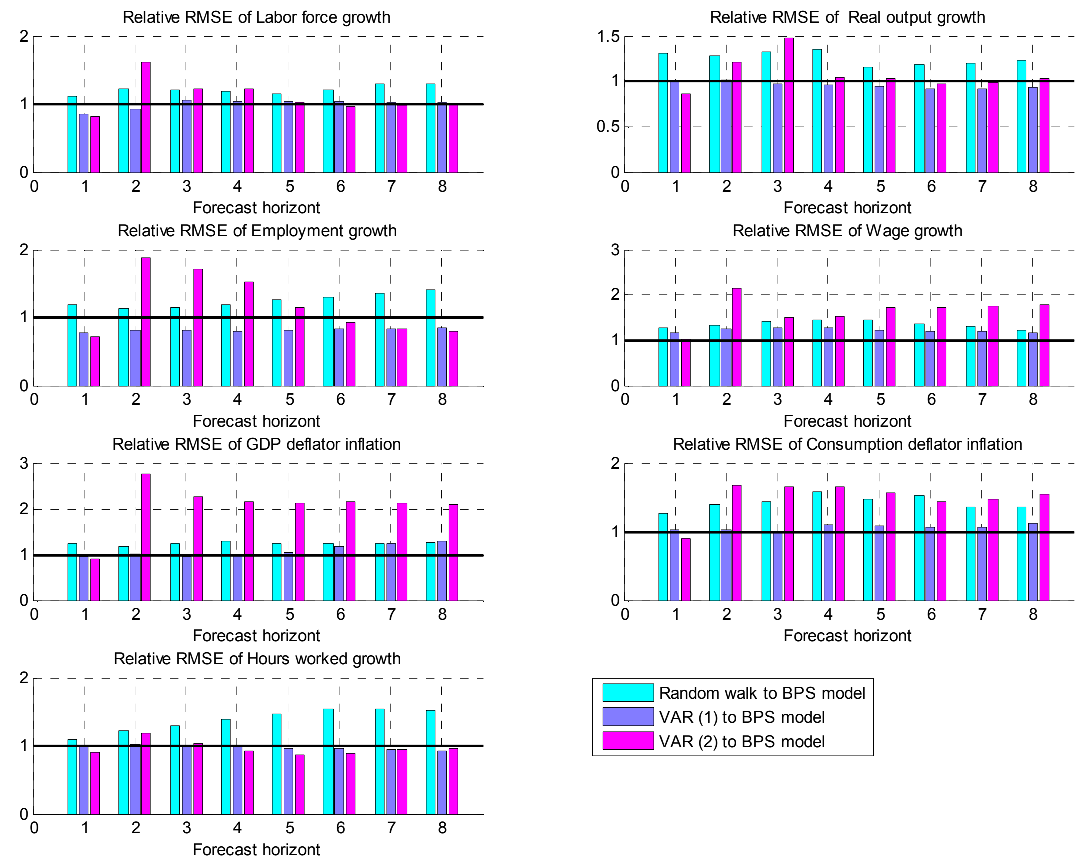
DOI: 10.7172/2353-6845.jbfe.2014.2.2

Figure 9
Forecasts' competition: Germany



Source: Authors' calculation.

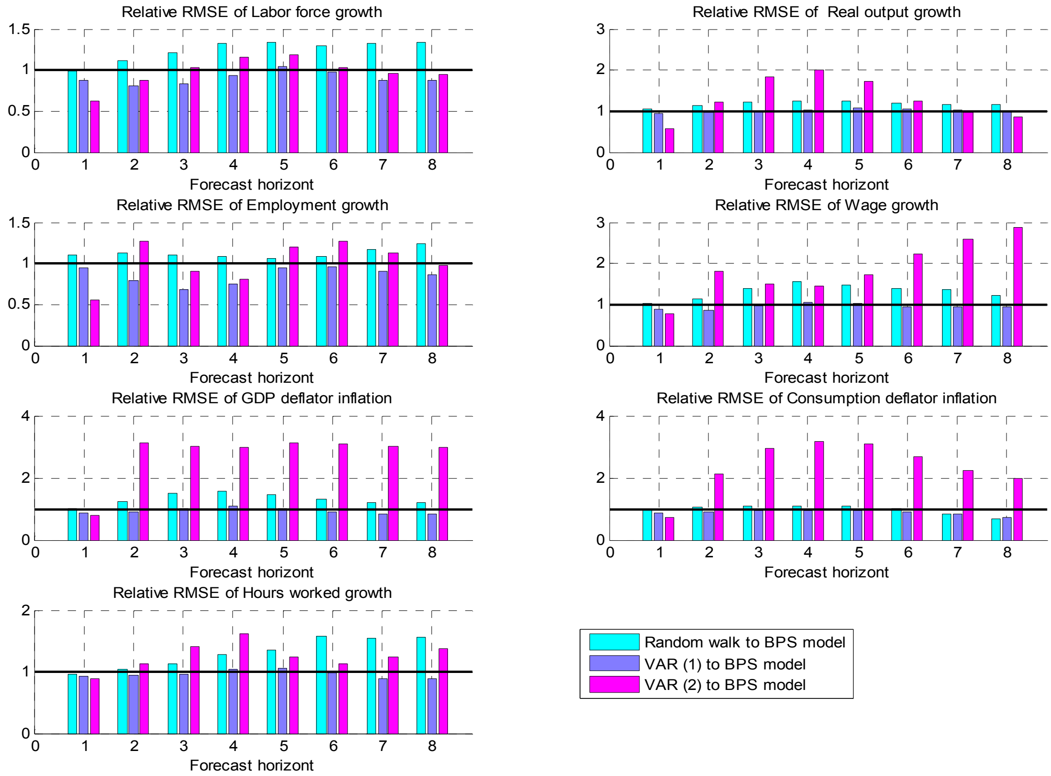
Figure 10
Forecasts' competition: France



Source: Authors' calculation.

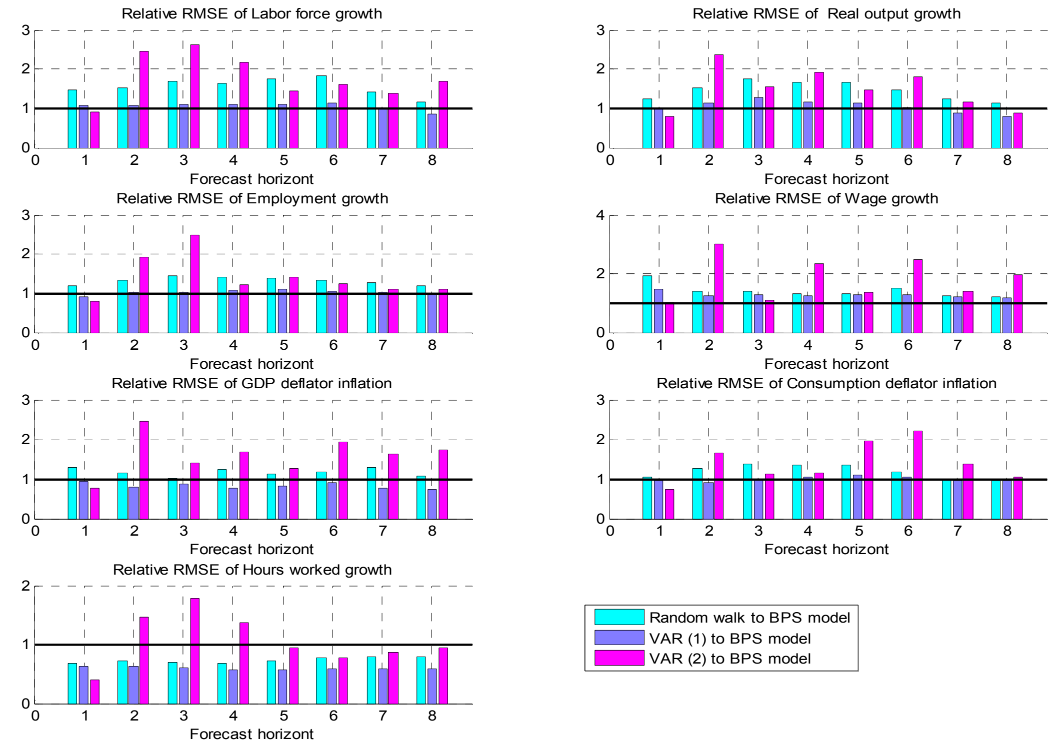
DOI: 10.7172/2353-6845.jbfe.2014.2.2

Figure 11
Forecasts' competition: Italy



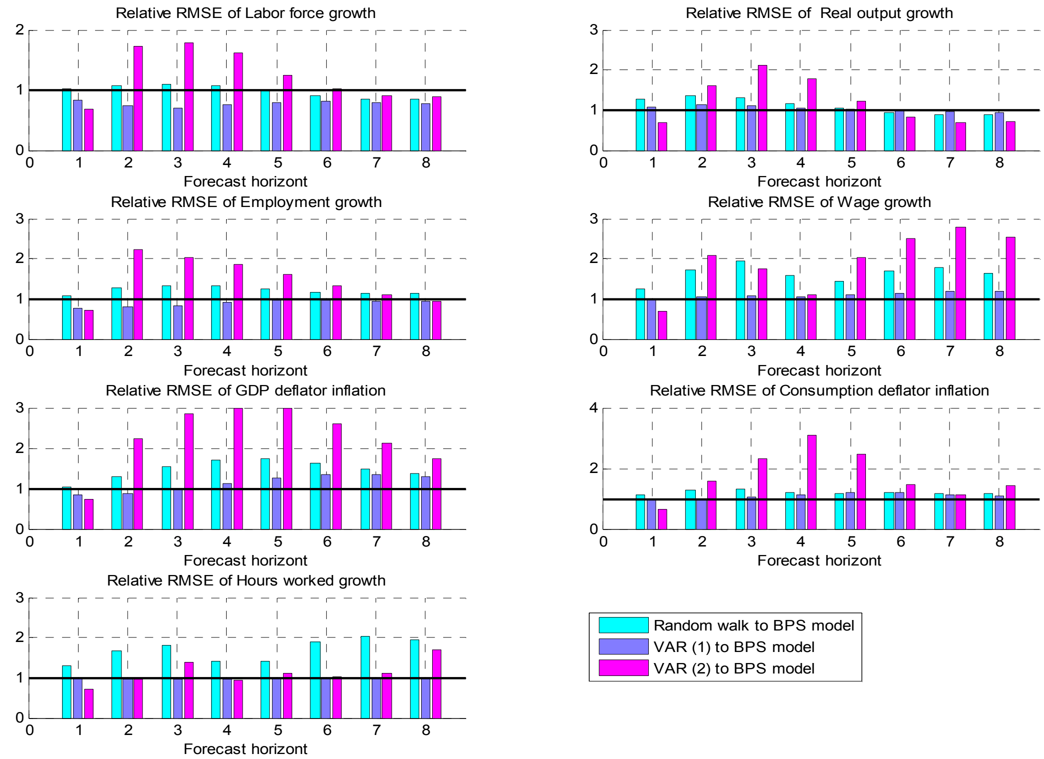
Source: Authors' calculation.

Figure 12
Forecasts' competition: Spain



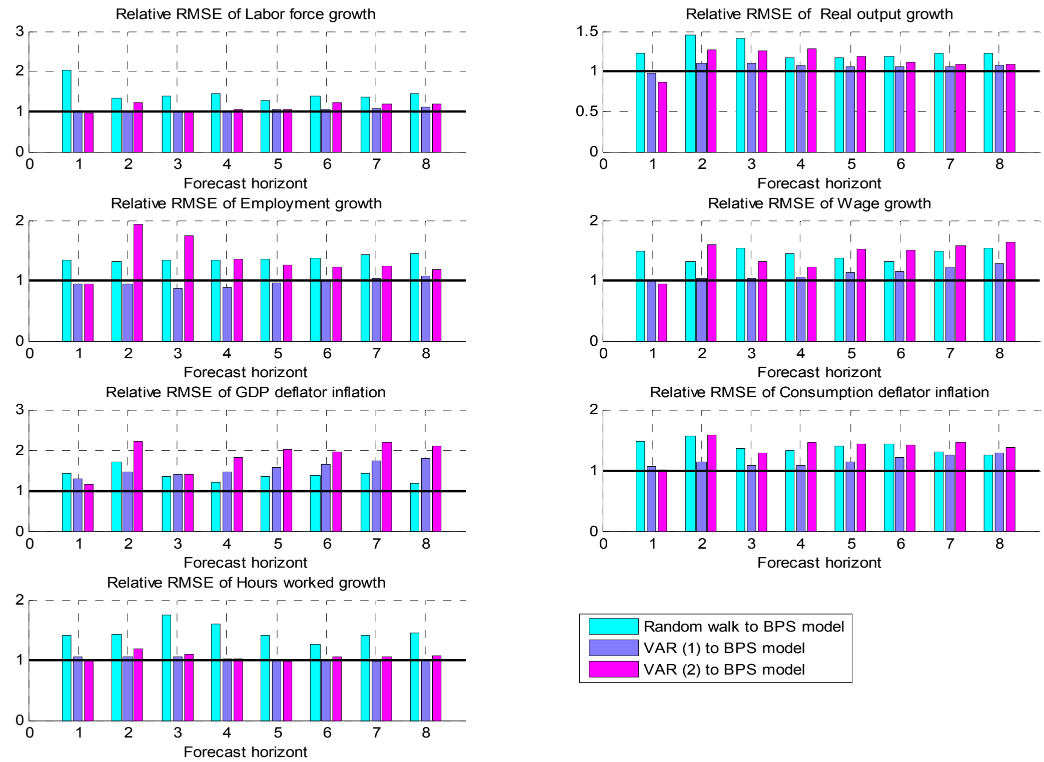
Source: Authors' calculation.

Figure 13
Forecasts' competition: The Netherlands



Source: Authors' calculation.

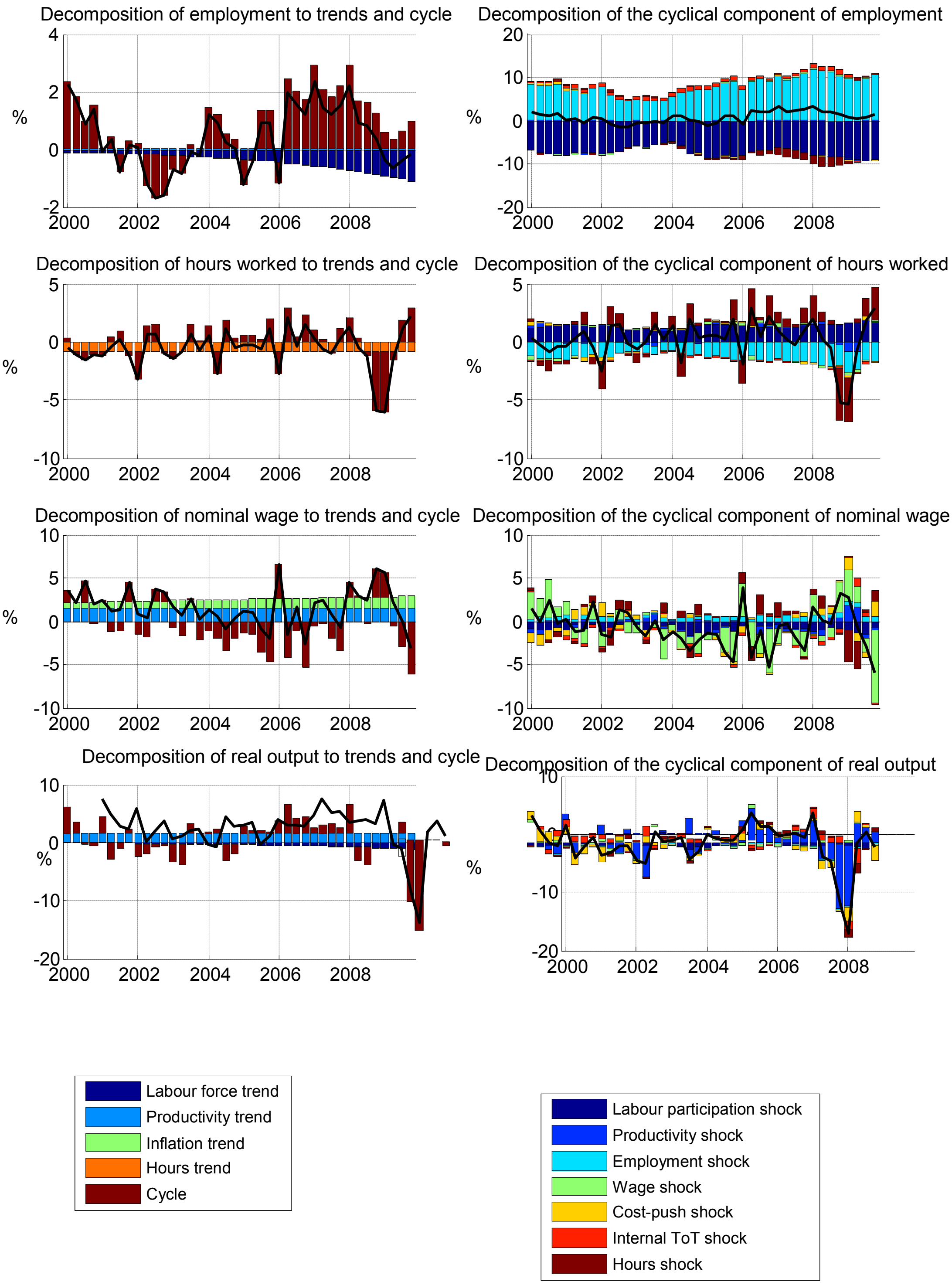
Figure 14
Forecasts' competition: Belgium



Source: Authors' calculation.

DOI: 10.7172/2353-6845.jbfe.2014.2.2

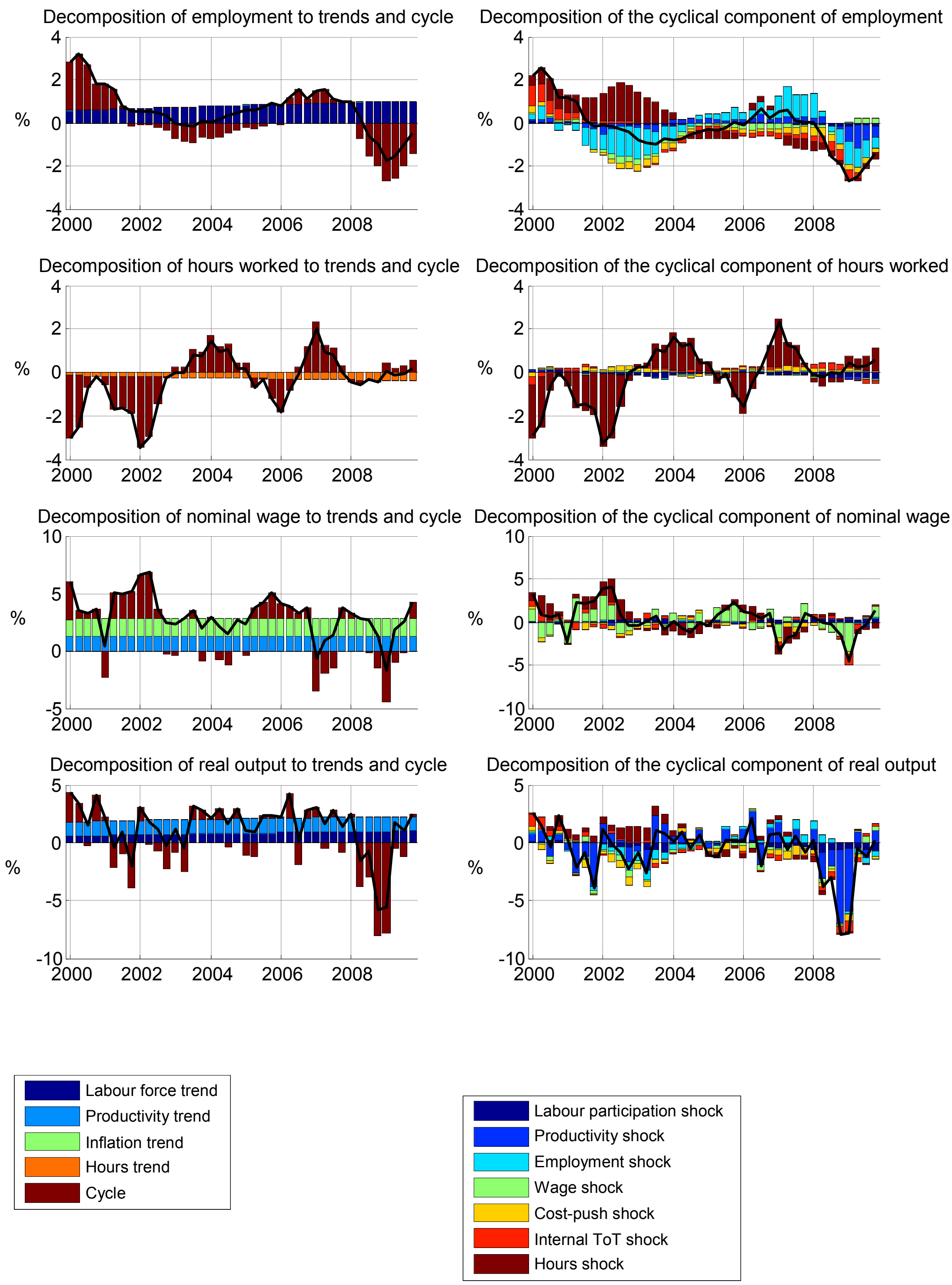
Figure 15
Historical decomposition for Germany



Note: figures reported in annualised growth rates.
Source: Authors' calculation.

DOI: 10.7172/2353-6845.jbfe.2014.2.2

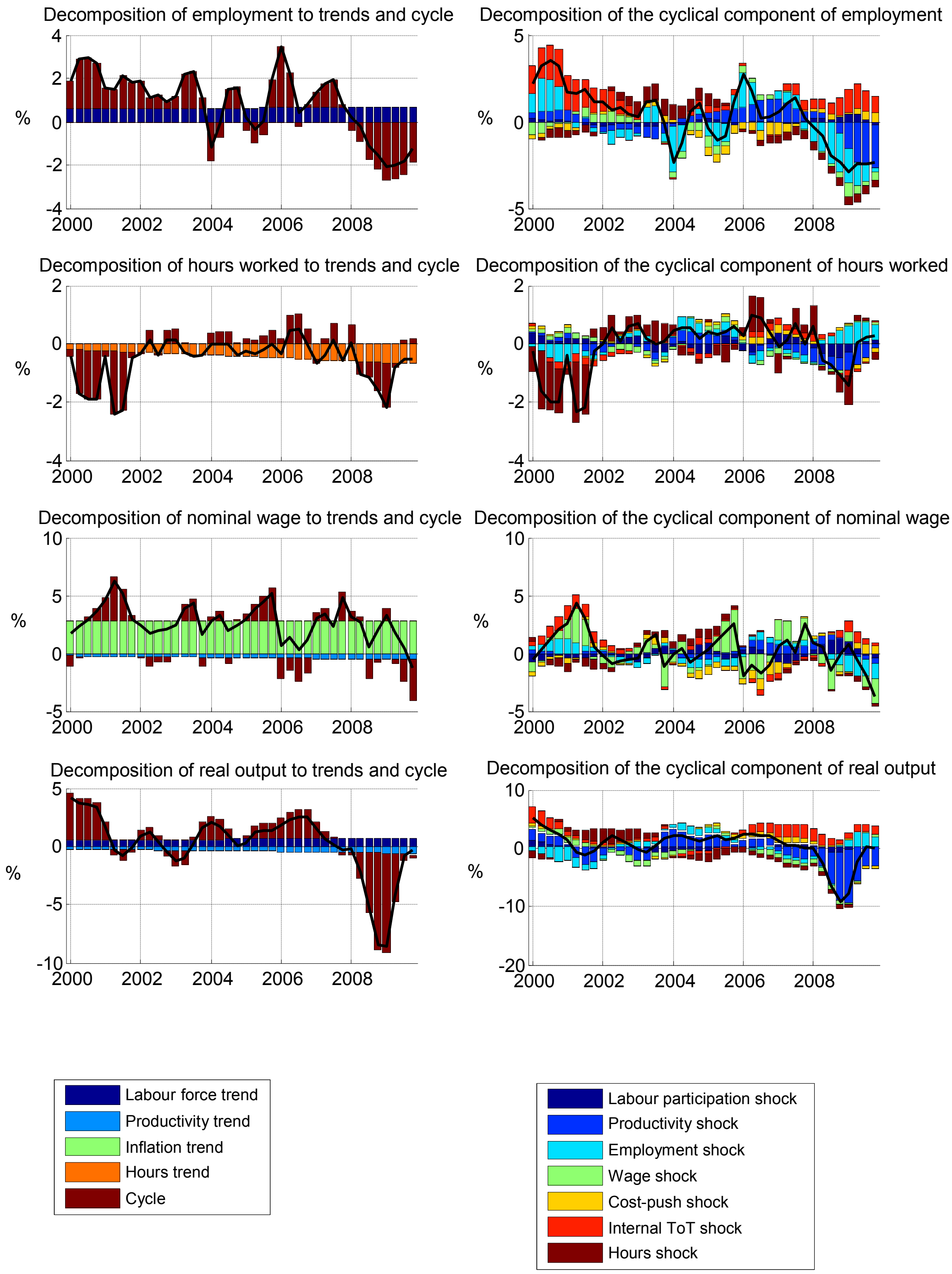
Figure 16
Historical decomposition for France



Note: figures reported in annualised growth rates.
Source: Authors' calculation.

DOI: 10.7172/2353-6845.jbfe.2014.2.2

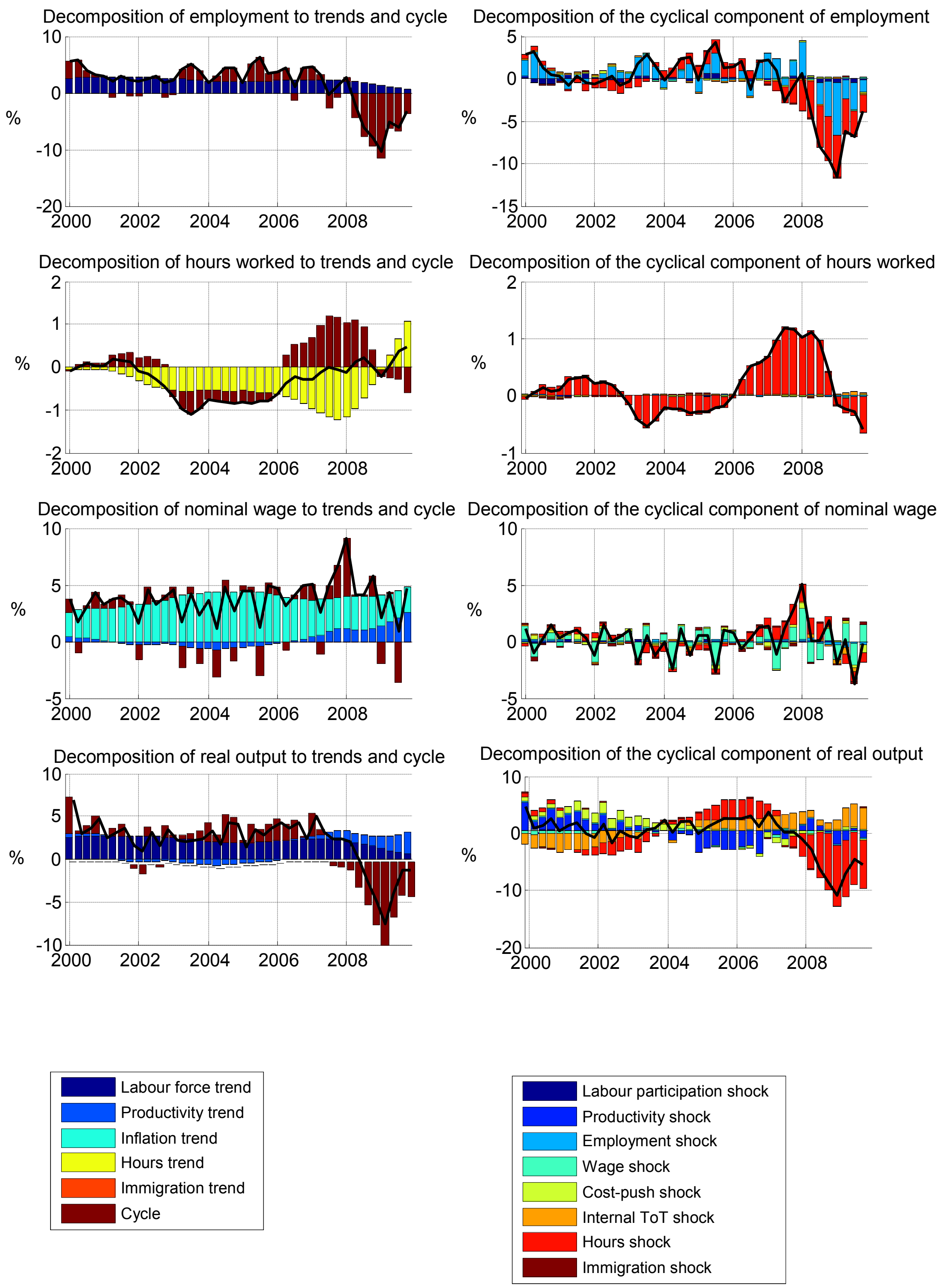
Figure 17
Historical decomposition for Italy



Note: figures reported in annualised growth rates.
Source: Authors' calculation.

DOI: 10.7172/2353-6845.jbfe.2014.2.2

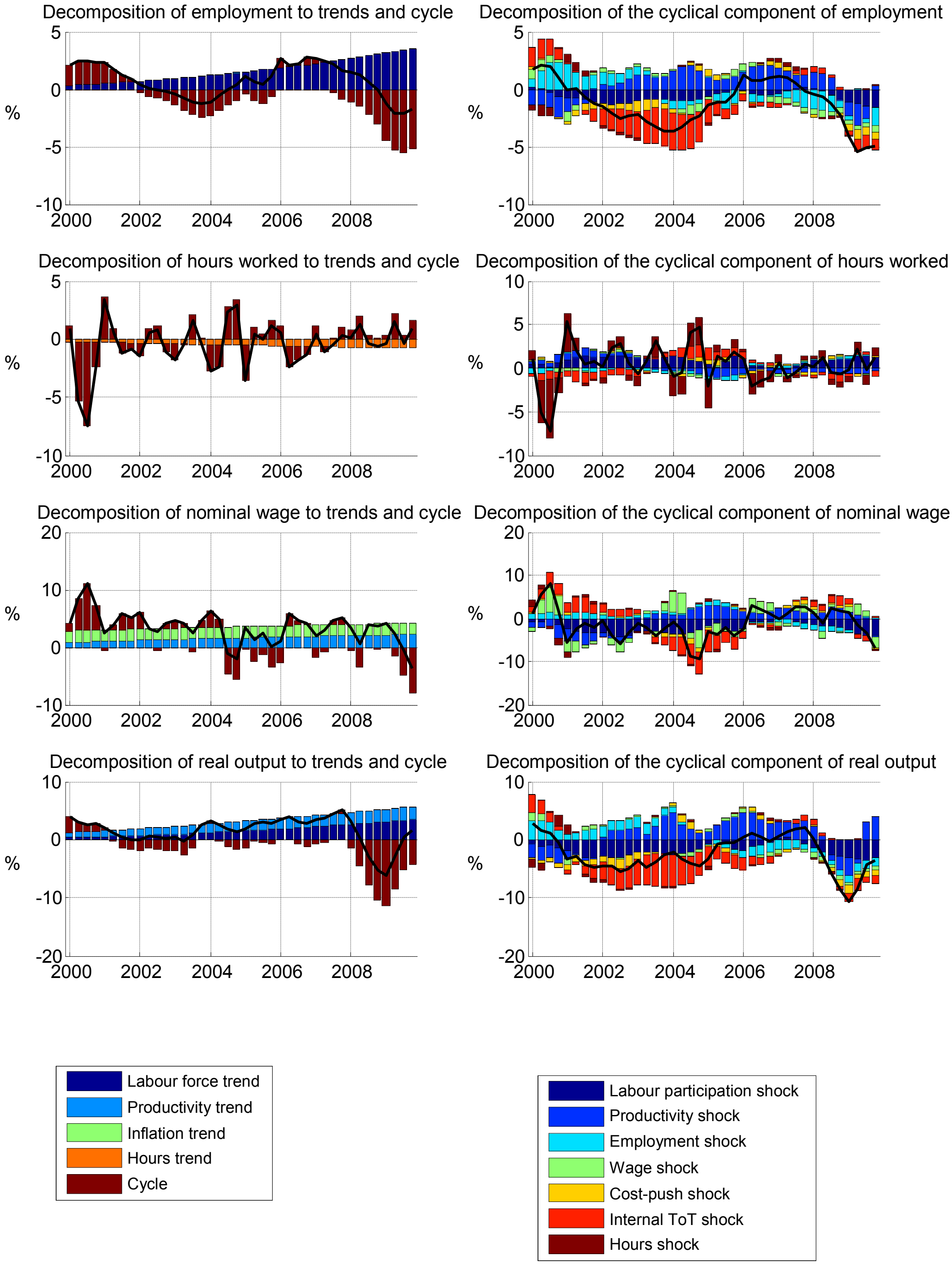
Figure 18
Historical decomposition for Spain



Note: figures reported in annualised growth rates.
Source: Authors' calculation.

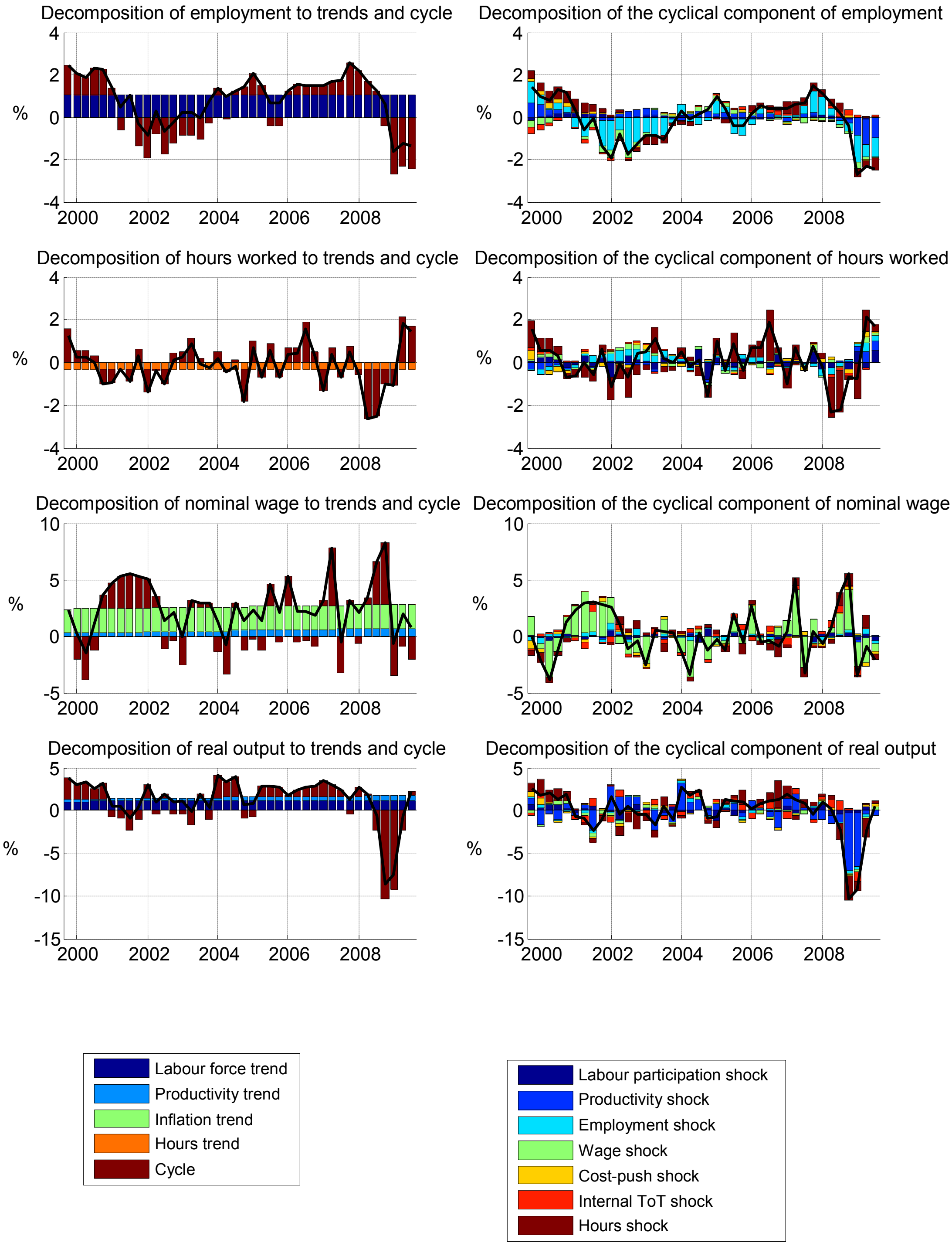
DOI: 10.7172/2353-6845.jbfe.2014.2.2

Figure 19
Historical decomposition for the Netherlands



Note: figures reported in annualised growth rates.
Source: Authors' calculation.

Figure 20
Historical decomposition for Belgium



Note: figures reported in annualised growth rates.
Source: Authors' calculation.

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Gravity Chains: Estimating bilateral trade flows when parts and components trade is important

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ABSTRACT

Trade is measured on a gross sales basis while GDP is measured on a net sales basis, i.e. value added. The rapid internationalisation of production in the last two decades has meant that gross trade flows are increasingly unrepresentative of the value-added flows. This fact has important implications for the estimation of the gravity equation. We present empirical evidence that the standard gravity equation performs poorly by some measures when it is applied to bilateral flows where the parts and components trade is important. We also provide a simple theoretical foundation for a modified gravity equation that is suited to explaining trade where international supply chains are important.

JEL classification: F01, F10

Keywords: Value chains, parts and components trade, gravity, bilateral flows

1. INTRODUCTION

Trade is measured on a gross sales basis wheras GDP is measured on a value-added basis. For the first decades of the postwar period, this distinction was relatively unimportant. Trade in intermediates was always important, but it was quite proportional to trade in final goods. The rapid internationalisation of supply chains in the last two decades has changed this (Yi 2003). Indeed, such trade has in recent decades boomed between advanced nations and emerging economies as well as among emerging nations – especially in Asia, where the phenomenon is known as ‘actory Asia’. There are, however, similar supply chains in Europe and between the US and Mexico (Kimura, Fukunari, Yuya Takahashi and Kazunobu Hayakawa 2007). As a result, gross trade flows are increasingly unrepresentative of the value-added flows. This fact has important policy implications (Lamy 2010), but it also has important implications for one of trade economists’ standard tools – the gravity equation.

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The basic point is simple. The standard gravity equation is derived from a consumer expenditure equation with the relative price eliminated using a general equilibrium constraint (Anderson 1979, Bergstrand 1985, 1989, 1990, Evenett and Keller, 2002). The corresponding econometrics widely used today is based on this theory (Anderson and Van Wincoop 2003, Anderson and Yotov, 2010, Novy 2010). As such the standard formulation – bilateral trade regressed on the two GDPs, bilateral distance and other controls – is best adapted to explaining trade in consumer goods. When consumer trade dominates, the GDP of the destination nation is a good proxy for the demand shifter in the consumer expenditure equation; the GDP of the origin nation is a good proxy of its total supply. By contrast, when international trade in intermediate goods dominates, the use of GDPs for the supply and demand proxies is less appropriate.

Consider, for instance, the determinants of Thai imports of auto parts from the Philippines. The standard formulation would use Thai GDP to explain Thailand's import demand, however, the underlying demand for parts is generated by Thai gross production of autos, not its value-added in autos. As long as the ratio of local to imported content does not change, value-added is a reasonable proxy for gross output, so the standard regression is likely to give reasonable results. However, for regions where production networks are emerging, value-added can be expected to be a poor proxy.

Why do incorrectly specified mass variables matter? A large number of gravity studies focus on variables that vary across country pairs – say free trade agreements, cultural ties, or immigrant networks. The most recent of these studies employ estimators that control for the mass variables with fixed effects. Such studies do not suffer from mass-variable mis-specification and so are unaffected by our critique. There are however a number of recent studies – especially concerning the 'distance puzzle' that stand proxy for the production and demand variables with GDP. It is these studies that our work speaks to.

For example, Rauch (1999), Brun et al. (2005), Berthelon and Freund (2008), and Jacks et al. (2008) use GDP as the mass variable when they decompose the change in the trade flow into the effects of income changes and trade cost changes; Anderson and Van Wincoop (2003) also use GDP as the mass variable in one of their estimation techniques. Since most of these studies are concerned with a broad set of nations and commodities, the mis-specification of the mass variable probably has a minor impact on the results – as the findings of Bergstrand and Egger (2010) showed. More worrying, however, is the use by authors that focus on trade in parts and components such as Athukorala and Yamashita (2006), Kimura et al. (2007), Yokota (2008), and Ando and Kimura (2009). These papers all use the consumer-good version of the gravity model to describe parts and components trade and thus have mis-specified the mass variable.

2. LITERATURE REVIEW AND PLAN OF THE PAPER

Literature Review

There is nothing new about trade in intermediates. Intermediate goods have long been important in the trade between the US and Canada; the 1965 US-Canada Auto Pact, for example, explicitly targeted preferential tariff reductions on cars and cars parts. It has also long been important within Western Europe as early studies of the EEC demonstrated (e.g. Dreze 1961, Verdoorn 1960, and Balassa 1965, 1966). A famous book by Grubel and Lloyd (1975), made clear that much of intra-industry trade was in intermediates, not final goods, and the importance of intermediates was reflected in early work by well-known theorists. For example, Vaneck (1963) presents an extension of the Heckscher-Ohlin model that allows for intermediates trade, and Ethier (1982) casts his model of intra-industry trade in a world where all trade was in intermediates.

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As better data and computing technology became available, the importance of intermediates in trade was rediscovered and documented more thoroughly. In the context of efforts to understand the impact of the EU's Single Market Programme, European scholars focused on the role of intermediates. For example, Greenaway and Milner (1987) list this as one of the 'unresolved issues', writing 'it is becoming increasingly obvious that a significant proportion of measured IIT is accounted for by trade in parts and components. [Nevertheless,] most of the models developed so far assume trade in final goods. The modelling of trade in intermediates needs to be explored further.' The issue attracted renewed interest following development of the new trade theory in the 1980s (Helpman and Krugman 1985) and again in the 1990s with Jones and Kierzkowski (1990), and Hummels, Rapoport and Yi (1998), and more recently Kimura, Takahashi and Hayakawa (2007), and Grossman and Rossi-Hansberg (2008).

The traditional gravity model was developed in the 1960s to explain factory-to-consumer trade (Tinbergen 1962, Poyhonen 1963, Linnemann 1966). This concept is at the heart of the first clear microfoundations of the gravity equation – a seminal paper by Anderson (1979). This article proposed a theoretical explanation of the gravity equation based on CES preferences when nations make a single differentiated product. Anderson and Van Wincoop (2003) use the Anderson (1979) theory to develop appropriate econometric techniques. Subsequent theoretical refinements have focused on showing that the gravity equation can be derived from many different theoretical frameworks (including monopolistic competition, and Melitz-type trade models with heterogeneous firms).

Studies on the gravity equations applicability to the intermediate-goods trade are more limited. These include Egger and Egger (2004), and Baldone et al. (2007). The study that is closest to ours is Bergstrand and Egger (2010). These authors develop a computable general equilibrium model that explains the bilateral flows of final goods, intermediate goods and FDI. Calibration and simulation of the model suggests a theoretical rationale for estimating a near-standard gravity model for the three types of bilateral flows. Using a large dataset on bilateral flows of final and intermediate goods trade, and a dataset on bilateral FDI flows, they estimate the three equations and find that the standard gravity variables all have the expected size and magnitude.

The value-added of our paper is primarily empirical – to show that the standard gravity specification performs poorly when applied to flows where trade in intermediates is important. Moreover, the failures line up with the predictions of our simple theory model that suggests a gravity equation formulation that is appropriate to intermediates trade. Note that when we perform the estimates on data pooled across a wide range of nations – as do Bergstrand and Egger (2010) – we find the same results, namely that the standard specification performs well. We believe the difference in the results is due to the fact that for many trade flows, the pattern of trade in intermediates is quite proportional to trade in final goods. This is especially true for trade among developed nations.

Plan of the paper

The paper starts with simple theory that generates a number of testable hypotheses. We then confront these hypotheses with the data and find that the estimated coefficients deviate from standard results in the way that the simple theory says they should. The key results are that the standard economic mass variable, which reflects consumer demand, does not perform well when it comes to bilateral trade flows where intermediates are dominant. Finally, we consider new proxies for the economic mass variables and show that using the wrong mass variable may bias estimates of other coefficients.

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3. THEORY

To introduce notation and fix ideas, we review the standard gravity derivation following Baldwin and Taglioni (2007).[2] Using the well-known CES preference structure for differentiated varieties, spending in nation-d ('d' for destination) on a variety produced in nation-o ('o' for origin) is:

$$v_{od} \equiv \left(\frac{p_{od}}{P_d}\right)^{1-\sigma} E_d; \qquad \sigma > 1 \tag{1}$$

where v_{od} is the expenditure in destination country-d, p_{od} is the consumer price inside nation-d of a variety made in nation-o, P_d is the nation-d CES price index of all varieties, s is the elasticity of substitution among varieties ($\sigma > 1$ is assumed throughout), and E_d is the nation-d consumer expenditure.

From the well-known profit maximization exercise of producers based in nation-o, $p_{od} = \mu_{od} m_o \tau_{od}$, where μ_{od} is the optimal price mark-up, m_o is the marginal costs, and τ_{od} is the bilateral trade cost factor, i.e. 1 plus the ad valorem tariff equivalent of all natural and manmade barriers. The mark-up is identical for all destinations if we assume perfect competition or Dixit-Stiglitz monopolistic competition; in these cases, the price variation is characterised by "mill pricing", i.e. 100% pass through of trade costs to consumers in the destination market.[3]

Here we work with Dixit-Stiglitz competition exclusively, so the mark-up is always $\sigma/(\sigma - 1)$. This means the local consumer price is $p_{oo} = (\sigma/(\sigma - 1))m_o\tau_{oo}$, where τ_{oo} is unity as we assume away internal trade barriers. Using this and summing over all varieties (assuming symmetry of varieties by origin nation for convenience), we have:

$$V_{od} = n_o p_{oo}^{1-\sigma} \frac{\tau_{od}^{1-\sigma}}{P_d^{1-\sigma}} E_d \tag{2}$$

where V_{od} is the aggregate value of the bilateral flow (measured in terms of the numeraire) from nation-o to nation-d; n_o is the number (mass) of nation-o varieties (all of which are sold in nation-d as per the well-known results of the Dixit-Stiglitz-Krugman model).

To turn this expenditure function (with optimal prices) into a gravity equation, we impose the market-clearing condition. Supply and demand match when (2) – summed across all destinations (including nation-o's sales to itself) – equals nation-o's output. When there is no international sourcing of parts, the nation's output is its GDP, denoted here as Y_o. Thus the market-clearing condition is: $Y_o = n_o p_{oo}^{1-\sigma} \sum_d \tau_{od}^{1-\sigma} P_d^{\sigma-1} E_d$. Solving this we obtain that $n_o p_{oo}^{1-\sigma} = Y_o / \Omega_o$ where W_o is the usual market-potential index (namely, the sum of partners' market sizes weighted by a distance-related weight that places lower weight on more remote destinations); specifically it is $\Omega_o \equiv \sum_d \tau_{od}^{1-\sigma} P_d^{\sigma-1} E_d$. Plugging this into (2) yields the traditional gravity equation:

$$V_{od} = \tau_{od}^{1-\sigma} E_d Y_o \frac{1}{P_d^{1-\sigma}} \frac{1}{\Omega_o} \tag{3}$$

Here P_d is the nation-d CES price index, while W_o is the nation-o market-potential index. It has become common to label the product $P_d^{1-\sigma}\Omega_o$ as the 'multilateral trade resistance' term. However, it is insightful to keep in mind the fact that 'multilateral trade resistance' is a combination of two well-known, well-understood, and frequently measured components.

[2] Another well-known derivation is from Helpman and Krugman (1985); they start from (1) and make supply-side assumptions that turns p_o into a constant, but makes n_{od} proportional to nation-o's GDP so the resulting gravity equation is similar – at least in the case of frictionless trade (the case they worked with in 1985).

[3] If one works with the Ottaviano Tabuchi and Thisse (2002) monopolistic competition framework, the mark-up varies bilaterally and so mill-pricing is not optimal.

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In the typical gravity estimation, E_d is proxied with nation-d's GDP, Y_d is proxied with nation-o's GDP, and t is proxied with bilateral distance.

3.1. Gravity when parts and components trade is important

To extend the gravity equation to allow for parts and components trade among firms, we need a trade model where intermediate-goods trade is explicitly addressed. It proves convenient to work with the Krugman and Venables (1996) 'vertical linkages'" model which focuses squarely on the role of intermediate goods. Here we present the basic assumptions and the manipulations that produce the modified gravity equation.

Krugman and Venables (1996) works with the standard new economic geography model where each nation has two sectors (a Walrasian sector, A, and a Dixit-Stiglitz monopolistic competition sector M), and a single primary factor, labour L. Production of A requires only L, but production of each variety of X requires L and a CES composite of all varieties as intermediate inputs (i.e. each variety is purchased both for final consumption and for use as an intermediate). Following Krugman and Venables (1996), the CES aggregate on the supply side is isomorphic to the standard CES consumption aggregate.

The indirect utility function for the typical consumer is:

$$V = I/P^c; \qquad P^c \equiv p_A^{1-\alpha}(P)^{\alpha}; \qquad P \equiv \left(\int_{i\in G} p_i^{1-\sigma} di\right)^{1/(1-\sigma)} \tag{4}$$

where I is consumer income, P^c is the ideal consumer price index, p_A is the price of A, the parameter "a" is the Cobb-Douglas expenditure share for M-sector goods, σ is the elasticity of substitution among varieties, P is the CES price index for M varieties, p_i is the consumer price of variety i, and G is the set of varieties available.

The cost function of a typical firm in a typical country is:

$$C[w,P,x] = (F + a_X x) w_{1-\alpha} P_{\alpha} \tag{5}$$

Here x is the output of a typical variety, F and a_X are cost parameters, w is the wage, and α is the Cobb-Douglas cost share for intermediate inputs.[4]

As noted above, mill pricing is optimal under Dixit-Stiglitz monopolistic competition. This, combined with the identity of the elasticity of substitution, σ, for each good's use in consumption and production, tells us that the price of each variety will be identical across the two types of customers. Choosing units such that $a_X = 1 - 1/\sigma$, the landed price will be:

$$p_{od} = \tau_{od} w_o^{1-\alpha} P_o^{\alpha}; \qquad \forall o,d \tag{6}$$

Using Shepard's and Hotelling's lemmas on (4) and (5), and adding the total demand for purchasers located in nation-d, we have an expression that is isomorphic to (2) except the definition of E now includes purchases by customers using the goods as intermediates:

$$V_{od} = n_o p_{oo}^{1-\sigma} \frac{\tau_{od}^{1-\sigma}}{P_d^{1-\sigma}} E_d; \qquad E_d \equiv \alpha(I_d + n_d C_d) \tag{7}$$

where I_d is nation-d's consumer income and C_d is the total cost of a typical nation-d variety.

[4] The assumption that the Cobb-Douglas parameter is identical in the consumer and producer CES price index is one of the strategic implications in the Krugman-Venables model; see their book for a careful examination of what happens when this is relaxed (Fujitu, Krugman and Venables 1999). The standard conclusion is that it does not qualitatively change results but it does significantly complicate the analysis in a way that requires numerical simulation.

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As before, we solve for the endogenous $n_o p_{oo}^{1-\sigma}$ using the market-clearing condition. In this case, the value that nation-o must sell is the full value of its M-sector output (not just its value-added). Under monopolistic competition's free entry assumption, the value of sales equals the value of full costs, so the market clearing equation becomes:

$$n_o C_o = n_o p_{oo}^{1-\sigma} \sum_d \tau_{od}^{1-\sigma} P_d^{1-\sigma} E_d; \qquad C_o \equiv C[w_o, P_o, x_o] \tag{8}$$

where the cost function C is given in (5). Solving (8) and plugging the result into (7) yields a gravity equation modified to allow for intermediates goods trade, namely:

$$V_{od} = \tau_{od}^{1-\sigma} E_d C_o \frac{1}{P_d^{1-\sigma}} \frac{1}{\Omega_o} \tag{9}$$

where E_d is defined in (7) and C_o is defined in (8), and $\Omega_o = \sum_d \tau_{od}^{1-\sigma} P_d^{\sigma-1} E_d$.

Expression (9) is the gravity equation modified to allow for trade intermediates. The key differences show up in the definition of the economic 'mass' variables since purchases are now driven both by consumer demand (for which income is the demand shifter) and intermediate demand (for which total production costs is the demand shifter).

4. BREAKDOWN OF THE STANDARD GRAVITY MODEL

This theoretical exercise suggests a key difference that should arise between gravity estimates on nations and time periods where most imports are consumer goods versus those where intermediates trade is important. Specifically, the standard practice of using the GDP of origin and destination countries as the 'mass' variables in the gravity equations is inappropriate for bilateral flows where parts and components are important. Of course, if the consumer- and producer-demand moves in synch – as they may in a steady-state situation – then GDP may be a reasonable proxy for both consumer and producer demand shifter. But if the role of vertical specialisation trade is changing over time, GDP should be less good at proxy-ing for the underlying demand shifters. For this reason, we expect that origin-country's GDP and destination country's GDP will have diminished explanatory power for those countries where value-chain trade is important.

These observations generate a number of testable hypotheses.

- The estimated coefficient on the GDPs should be lower for nations where the parts and components trade is important, and should fall as the importance of parts trade rises.
- As vertical specialisation trade has become more important over time, the GDP point estimates should be lower for more recent years.
- In those cases where the GDPs of the trade partners lose explanatory power, bilateral trade should be increasingly well explained by demand in third countries.

For example, China's imports should shift from being explained by China's GDP to being explained by its exports to, say, the US and the EU. There are two ways of phrasing this hypothesis. First, China's imports are a function of its exports rather than its own GDP. Second, China's imports are a function of US and EU GDP rather than its own, since US and EU GDP are critical determinants of their imports from China.

To check these conjectures, we estimate the standard gravity model for different sets of countries and sectors for a panel that spans the years 1967 to 2007. We run standard log-linear gravity equations using pooled cross-section time series data, namely:

$$\ln(V_{odt}) = G + \alpha_1 \ln\left(\frac{Y_{ot}}{\Omega_{ot}} \cdot \frac{E_{dt}}{P_{dt}^{1-\sigma}}\right) + \alpha_2 \ln \tau_{odt} + \varepsilon_{odt} \tag{10}$$

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A key econometric problem is that the price index P_{dt} and the market potential index W_{ot} are unobservable and yet include factors that enter the regressions independently (e.g. E, Y and t). Thus ignoring them can lead to serious biases.

If the econometrician is only interested in estimating the impact of a pair-specific variable – such as distance or tariffs – the standard solution is to put in time-varying country-specific fixed effects. This eliminates all the terms multiplied by a_1 in equation (10). Plainly we cannot use this approach to investigate the impact of using GDPs as the economic mass proxies when trade in parts and components is important. We thus need other means of controlling for Ω_{ot} and P_{dt}.

Our baseline specification accounts for the terms Ω_{ot} and P_{dt} explicitly. As precise measures of Ω_{ot} and P_{dt} are hard to construct, we perform robustness checks using fixed effects specifications. To ensure comparability with the fixed effects specification, in the key specifications we enter the importer's and exporter's economic mass as a single product-term into the equation, with the shortcoming of forcing the coefficient of the importer and exporter mass variables to be the same. Specifically, the term accounting for the product of the trade partners' economic mass is the product of importer-d real GDP (so to account for P_{dt}) and of exporter-o's nominal GDP divided by a proxied for Ω_{ot}, constructed adapting a method first introduced by Baier and Bergstrand (2001) namely:

$$\Omega_{ot} = \left(\sum\nolimits_d GDP_{dt} * (Dist_{od})^{1-\sigma}\right)^{\frac{1}{1-\sigma}}$$

The elasticity value in the Ω_{ot} relationship has been set as $\sigma = 4$, which corresponds to estimates proposed in empirical literature (e.g. Obstfeld and Rogoff 2001 and Carrere 2006).

Turning to the trade cost variable, t, we introduce standard trade frictions, including log of bilateral distance, and dummies for contiguity, and common language. Moreover for robustness purposes we also test for additional time-varying trade frictions measured by *cif-fob* ratios, as proposed by Bergstrand and Egger (2010).

The data used for the bilateral trade flows, and the cif-fob ratios are taken from the UN COMTRADE database. GDPs are from the World Bank's World Development Indicators. Bilateral distances, contiguity, and common language are from the CEPII database. Data for Taiwan, which are missing from the UN databases, are from CHELEM (CEPII) and national accounts.

Estimation is by simple ordinary least squares with the standard errors clustered by bilateral pairs since we work in direction-specific trade flows rather than the more traditional average of bilateral flows.

4.1. Empirical results

In Table 1 we report the gravity equation estimates for all goods as well as for intermediate and final goods separately. Intermediate and final goods have been identified according to the UN Broad Economic Categories Classification (see appendix). The sample includes all the countries where data is available, namely 187 nations.

Coefficients have the expected signs and are statistically significant. For all six regressions (all goods, only intermediates, and only consumer goods with and without time fixed effects) the estimates are broadly similar. The mass variables are all estimated to be close to unity. The bilateral distance variable is negative and falls in the expected range. The additional trade cost measure, the cif/fob ratio, is always negative as expected. Continuity and language always have the expected sign and fall in the usual ranges.

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Table 1
Bilateral flows of total, intermediate and final goods, 187 nations, 2000–2007

Variables	All goods		Intermediates only		Consumer goods only	
	(1)	(2)	(3)	(4)	(5)	(6)
ln $(GDPo_t{}^*GDPd_t/\Omega ot^*Pdt)$	0.860***	0.865***	0.898***	0.905***	0.791***	0.796***
	(0.006)	(0.006)	(0.007)	(0.007)	(0.008)	(0.008)
ln(cif/fob ratio)	-0.083***	-0.080***	-0.189***	-0.184***	-0.341***	-0.338***
	(0.013)	(0.013)	(0.015)	(0.015)	(0.017)	(0.017)
ln Distance	-0.775***	-0.777***	-0.851***	-0.855***	-0.758***	-0.760***
	(0.019)	(0.019)	(0.022)	(0.022)	(0.025)	(0.025)
Contiguity	1.575***	1.565***	1.711***	1.697***	1.356***	1.347***
	(0.105)	(0.105)	(0.119)	(0.119)	(0.127)	(0.127)
Common language	0.966***	0.972***	0.997***	1.005***	1.186***	1.192***
	(0.046)	(0.046)	(0.052)	(0.052)	(0.059)	(0.059)
Constant	-28.61***	-28.74***	-30.84***	-31.03***	-26.87***	-27.02***
	(0.359)	(0.363)	(0.400)	(0.404)	(0.456)	(0.459)
Time dummies		yes		yes		yes
Observations	62875	62875	62875	62875	58468	58468
R-squared	0.627	0.628	0.585	0.587	0.479	0.480

Note: Dependent variable: imports + re-imports. Standard errors are clustered by bilateral pair. Robust standard errors are reported in parenthesis: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Source: Authors' calculations.

These Table 1 results confirm the findings of Bergstrand and Egger (2010), namely that the size of the estimated coefficients does not vary for consumer and intermediate goods. As such, it would seem that our concern about mis-estimating the gravity equation is misplaced. However, as noted above, if the consumer and intermediate trade is roughly proportional over time, GDP will be a reasonable proxy for both consumer income and gross value-added. The real test of the stability of the parameters would be on a sample where the importance of the intermediates trade was rising significantly.

To check this, we turn to a sub-sample of nations where we a priori expect intermediate trade to be both very important and growing more rapidly than consumer trade. Specifically, we estimate a gravity model as in Table 1, but on bilateral trade between pairs of 'Factory Asia' countries (i.e. Japan, Indonesia, Korea, Malaysia, Philippines, Thailand, and Taiwan). To gauge the stability of parameters, we interact period-dummies with the mass variable. The results, shown in Table 2, are quite different to those of Bergstrand and Egger (2010) and to those of Table 1.

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Table 2
Bilateral flows of total goods among 'Factory Asia' nations (1967–2008)

Variables	**No time interactions**			**Variable mass coefficient**	
	(1)	**(2)**	**(3)**	**(4)**	**(5)**
ln ($GDP_o \times GDP_d / \Omega_o \times P_d$)	0.725***	0.725***	0.764***	0.425***	0.504***
	(0.009)	(0.028)	(0.026)	(0.055)	(0.051)
*years 1967–1986				0.318***	0.278***
				(0.048)	(0.048)
*years 1987–1996				0.177***	0.164***
				(0.027)	(0.032)
*years 1998–2002				0.007	0.00274
				(0.015)	(0.017)
ln (Distance)	-0.258***	-0.258		-0.0414	
	(0.0570)	(0.298)		(0.297)	
Contiguity	0.188***	0.188		0.167	
	(0.0682)	(0.386)		(0.367)	
Colony	-0.487***	-0.487		0.0695	
	(0.101)	(0.388)		(0.405)	
Common coloniser	-0.620***	-0.620*		-0.296	
	(0.116)	(0.325)		(0.324)	
Constant	-7.218***	-7.218***	-8.825***	-1.465	-2.632**
	(0.433)	(2.281)	(0.485)	(2.279)	(1.178)
Time effects	yes	yes			
Exporter*time effects			yes	yes	Yes
Importer*time effects			yes	yes	Yes
Pair effects			yes	yes	Yes
Clustered Standard Errors		yes	yes	yes	Yes
Observations	1722	1722	1722	1722	1722
R-squared	0.833	0.833	0.936	0.851	0.948

Note: Standard errors are clustered by bilateral pair. Robust standard errors in parenthesis: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. 'Factory Asia' countries: Japan, Indonesia, Republic of Korea, Malaysia, Thailand, and Taiwan.

Source: Authors' calculations.

The baseline regressions (without time interactions) show the fairly common result that the gravity model does not work well on 'Factory Asia' nations. The estimated mass coefficient is fairly low at about 0.7. The distance estimate, however, at -0.26 is much lower than the commonly observed -0.7 to -1.0. When we include time-interaction terms for the economic mass variable, we find that the coefficient is not stable over time. When the standard controls are included, see column (4), the base case estimate is 0.4 to which must be added the period coefficients which are 0.3 for the pre-Factory Asia period (Baldwin 2006), 0.2 for the 1987–1996 period, and essentially zero (and insignificant) for the post 1998 period.

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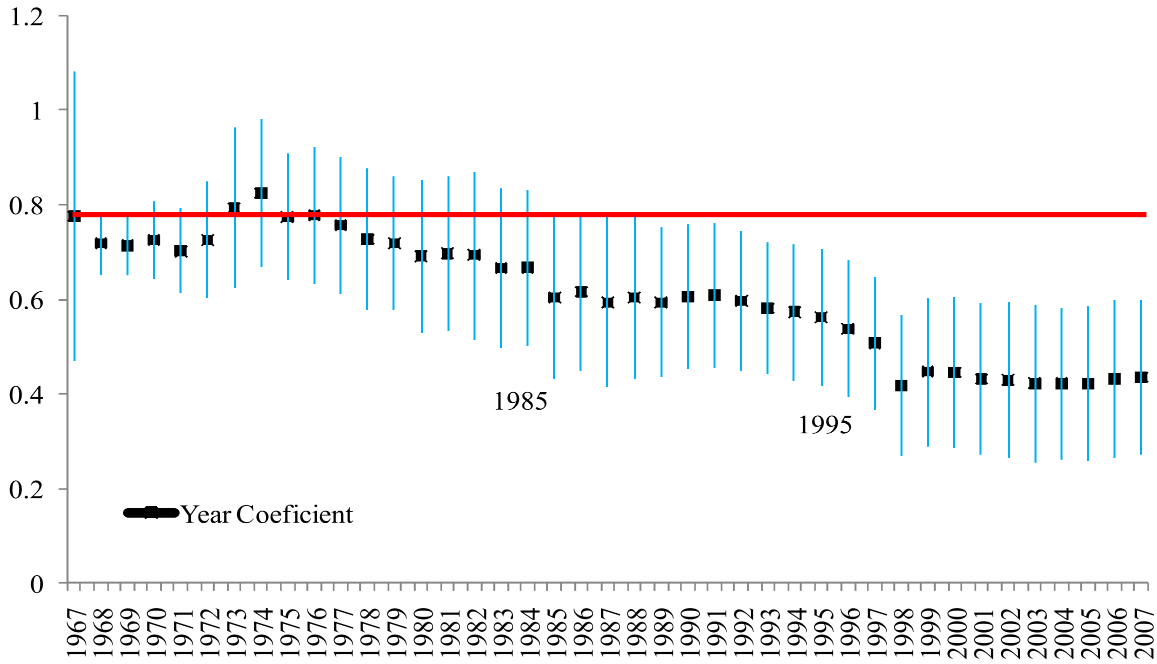
Figure 1
GDP coefficients for Factory Asia countries, 1967–2008



Notes: Estimated mass-elasticity coefficients with year interactions and pair fixed effects (as in (10). High and low bars show plus/minus 2 standard errors; Factory Asia countries: Japan, Indonesia, Republic of Korea, Malaysia, Thailand, and Taiwan.

To estimate the mass variable's instability over time more clearly, we re-do the same regression but allowing yearly interaction terms. The results, displayed in Figure 1, show the evolution of the GDP coefficients. The mass elasticity fall over time, with two clear breaks in the estimated coefficients, 1985 and 1998.

The timing and direction of these structural changes are very much in line with the literature on the internationalisation of production. According to many studies, production unbundling started in the mid-1980s and accelerated in the 1990s (e.g. Hummels, Rapport and Yi 1998). The idea is that coordination costs fell with the ICT revolution and this permitted the spatial bundling of production stages (Baldwin 2006). The ICT revolution came in two phases. The internet came online in a massive way in the mid-1980s, and then, in the 1990s, the price of telecommunications plummeted with various ITC-related technical innovations and widespread deregulation (Baldwin 2011). The upshot of all these changes was that it became increasingly economical to separate manufacturing stages geographically. Stages of production that previously were performed within walking distance to facilitate face-to-face coordination could be dispersed without an enormous drop in efficiency or timeliness.

As far as the Figure 1 results are concerned, the notion is that as trade became increasingly focused on intermediates, GDP became an increasingly poor determinant of trade flows – as suggested by our theory. The impact of the mid-1980s changes and the mid-1990s changes are clear from the estimated GDP elasticities. More specifically, from 1967 to 1985 the elasticity of these countries' bilateral imports to GDP was stable, with a coefficient of about 0.77. Between 1985 and 1997, it steadily decreased to reach a coefficient value of about 0.60, and after 1998, it further dropped to a figure close to 0.40. The coefficient estimates for the different periods in Factory Asia are summarised in Table 2 , columns (4) and (5).

For sake of comparison we also report results of time-year interactions with GDP for bilateral trade between countries where we a priori expect bilateral trade to be dominated by consumption

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goods and/or a stable ratio of intermediates to the final goods trade. To this end, we re-run the Table 2 regressions for exports and imports by the United States and Australia, with the EU15 nations as trade partner. Because most of the internationalisation of supply chains is regional rather than global (except for microelectronics), we expect these bilateral trade flows to be less influenced by the second unbundling that so marked Factory Asia trade. The results, shown in Table 3 tend to confirm our view that the gravity model breaks down only for bilateral flows where production sharing is especially important and growing quickly. That is, as predicted by our theory, we find no breaks over time in the trade coefficients while distance coefficients have elasticity levels which are closer to unity. None of the time interaction terms in columns (4) and (5) are significant and the other point estimates fall in the expected ranges.

Table 3
Estimates for EU15, and US, Canada, Australia and New Zealand, 1967–2008

Variables	No time interactions			Variable mass coefficient	
	(1)	(2)	(3)	(4)	(5)
$\ln(GDP_o \times GDP_d / \Omega_o \times P_d)$	0.659***	0.659***	0.632***	0.725***	0.703***
	(0.009)	(0.025)	(0.027)	(0.058)	(0.034)
*years 1967–1986				-0.0408	-0.0503
				(0.051)	(0.044)
*years 1987–1996				-0.0376	-0.0444
				(0.036)	(0.032)
*years 1998–2002				0.0132	0.005
				(0.017)	(0.014)
ln (Distance)	-0.843***	-0.843***		-0.688**	
	(0.059)	(0.233)		(0.276)	
Constant	-1.630**	-1.630	-8.819***	-4.966	-10.72***
	(0.726)	(2.284)	(0.657)	(3.733)	(0.917)
Time effects	yes	yes			
Exporter*time effects			yes	yes	yes
Importer*time effects			yes	yes	yes
Pair effects			yes	yes	yes
Observations	820	820	820	820	820
R-squared	0.932	0.932	0.978	0.934	0.978
Clustered Standard Errors		yes	yes	yes	yes

Note: Standard errors are clustered by bilateral pair. Robust standard errors are reported in parenthesis: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.
Source: Authors' calculations.

4.2. More precise estimates of the impact of components on the mass estimate

These two sets of results are highly suggestive. On data that is widely recognised as being dominated by parts and components trade, we find structural instability in the mass variable coefficient moving in the expected direction. However, on data where this sort of production fragmentation is not widely viewed as having been important, we find that that mass point-estimates are stable over time.

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To explore this more systematically, we consider a more continuous relationship between the importance of components trade and the point-estimate on the mass variable on the full sample. Our basic assertion is that the composition of trade flows will influence the point estimates of the economic mass variables since the standard gravity model is mis-specified when it comes to the mass variable. The most direct test of this hypothesis is to include the ratio of intermediates to total trade as a regressor, both on its own and – more importantly – as an interaction term with the economic mass variable. Of course a mis-specification of one part of the regression has implications for the point-estimates of the other regressors, so we also consider the ratio's interaction with the other main regressors.

To this end, we re-estimate the basic equation on the full sample of 187 countries for the years 2000-2008 allowing for interactions with a variable that accounts for the share of intermediate goods over total imports in each particular bilateral trade flow.

The idea here is that GDP as a measure for economic mass should work less well for those bilateral flows that are marked by relatively high shares of intermediates trade. By estimating the effect on the full sample, we avoid the problem of identifying the exact sources of the variation in the coefficients. We implement the idea in two ways.

First we estimate the standard regression but include the share of bilateral imports that is in intermediates (denoted as M_d^{interm}/M_d). This new variable is included on its own and interacted with the other right-hand side variables. Table 4 reports the estimated results for the coefficients of interest.

The regression results tend to confirm our hypothesis. The regression reported in column (1) includes the ratio on its own and interacted only with the mass variable. The coefficients for economic mass and distance are a very reasonable at 1.031 and -1.173 respectively (both significant at the 1% level). The ratio on its own comes in positive as expected (bilateral trade-links marked by a high share of intermediates tend to have 'too much' trade compared to the prediction of the standard gravity equation). The ratio interacted with economic mass also has a negative sign, -0.129, which conforms with our hypothesis (the higher is the ratio of intermediates for the particular trade pair, the lower is the estimate of the economic mass variable). All coefficients are significantly different to zero at the 1% level of confidence.

The other columns report robustness checks on the main regression. The qualitative results on the variables of interest (the mass coefficient, the ratio coefficient, and the mass*ratio interaction coefficient) are robust to inclusion of interaction terms with any or all of the control variables. This confirms the more informal tests based on an a priori separation of the sample.

Interestingly, the interaction term is also highly significant and negative for distance in specification (2). That is, distance seems to matter more for parts and components trade – a result that is not in line with our simple model, but is expected from the broader literature on offshoring. For example, transportation costs become more important when trade costs are incurred between each stage of production while the value added per stage is modest.

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Table 4
Interactions with share of intermediates in total imports, full sample

Variables	**(1)**	**(2)**	**(3)**	**(4)**
M_d^{interm}/M_d	6.536***	8.018***	6.954***	7.330***
	(0.858)	(1.015)	(0.835)	(1.004)
ln $(GDP_o \times GDP_d/\Omega_o \times P_d)$	1.031***	1.027***	1.064***	1.058***
	(0.010)	(0.010)	(0.010)	(0.010)
* M_d^{interm}/M_d	-0.129***	-0.118***	-0.137***	-0.126***
	(0.017)	(0.017)	(0.017)	(0.016)
ln (Distance)	-1.173***	-1.051***	-1.011***	-0.954***
	(0.018)	(0.037)	(0.0191	(0.037)
* M_d^{interm}/M_d		-0.232***		-0.110*
		(0.059)		(0.0601
$Contig_{od}$			1.350***	0.967***
			(0.101)	(0.246)
* M_d^{interm}/M_d				0.625*
				(0.369)
Common language			1.215***	1.126***
			(0.044)	(0.078)
* M_d^{interm}/M_d				0.178
				(0.119)
Constant	-27.58***	-28.40***	-30.85***	-31.07***
	(0.551)	(0.634)	(0.541)	(0.625)
Observations	121737	121737	121737	121737
R-squared	0.604	0.604	0.621	0.621

Notes: $M_d^{tinterm}/M_d$ is the share of intermediate imports by a country d over its total imports. Robust standard errors are reported in parenthesis: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

The second approach is to use decile-dummies to permit a more flexible relationship between the share of imports made up of components and the mass point-estimate. The idea is that the inclusion of the intermediates-ratio imposes linearity on the relationship. The deciles approach allows the interaction terms to be non-linear, for example it allows for the possibility of a threshold effect whereby the interaction is significant but only for ratios that are sufficiently large. More specifically, the dummies categorise the share of intermediates in total imports, i.e. a dummy that selects bilateral flows where the proportion of intermediate imports is below 10%, between 10% and 20%, etc. The results are shown in Table 5. All results are robust to the addition of other trade determinants.

For the variable of greatest interest, the economic mass variable, the coefficient for the base-case decile is 0.985 which is very close to unity as expected and very precisely estimated. The subsequent rows show the additional effects for each decile. What we see is that the interaction terms are insignificant for shares of intermediates below 50% of total imports. However, for high concentrations of intermediates, the interaction terms are all negative and highly significant – at the 1% level. The additional effects lower the base case point-estimate by around 0.10. The distance term is a very reasonable -1.1 and highly significant.

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Table 5
All countries, 2000–2007, by share of intermediate imports

Variables	$(GDP_o \times GDP_d/\Omega_o \times P_d)$	ln(Distance)	Constant
Base effect	0.985***	-1.105***	-26.29***
	(0.018)	(0.018)	(0.898)
Base effect * d2	-0.0308		
	(0.021)		
Base effect * d3	0.0108		
	(0.021)		
Base effect * d4	-0.0330		
	(0.020)		
Base effect * d5	-0.0803***		
	(0.020)		
Base effect * d6	-0.103***		
	(0.021)		
Base effect * d7	-0.0903***		
	(0.021)		
Base effect * d8	-0.0723***		
	(0.022)		
Base effect * d9	-0.118***		
	(0.024)		
Base effect * d10	-0.0748***		
	(0.022)		
Observations	121712		
R-squared	0.610		

Note: deciles categorise countries' bilateral imports by increasing shares of intermediate imports over total imports. Hence q10 indicates the 10% bilateral import relationships where the share of intermediate imports in total imports is highest and the base effect the 10% bilateral import relationships where the share of intermediate imports in total imports is lowest. Common language and contiguity included but not reported. Standard errors are clustered by bilateral pair. Robust standard errors are reported in parenthesis: *** $p < 0.01$, ** $p < 0.05$, * $p < 0$.

Source: Authors' estimations.

The results in Table 5 suggests that there is something of a threshold effect in operation. What we see is that the standard gravity specification works rather well for bilateral trade flows where the ratio of intermediates is not too great. For trade flows where intermediates are more important, however, we get the by now familiar result that the mass coefficient is significantly lower. Since this share is indeed rather low for most bilateral trade flows in the world (since production fragmentation tends to be a regional phenomenon), this may help explain the Baier and Egger (2010) result mentioned above.

To illustrate the point graphically, we plot, in Figure 2, the point estimates and standard errors using a candle chart. Here the point estimates of the mass coefficients are plotted as the horizontal bar; the associated standard errors are show with the vertical bar.

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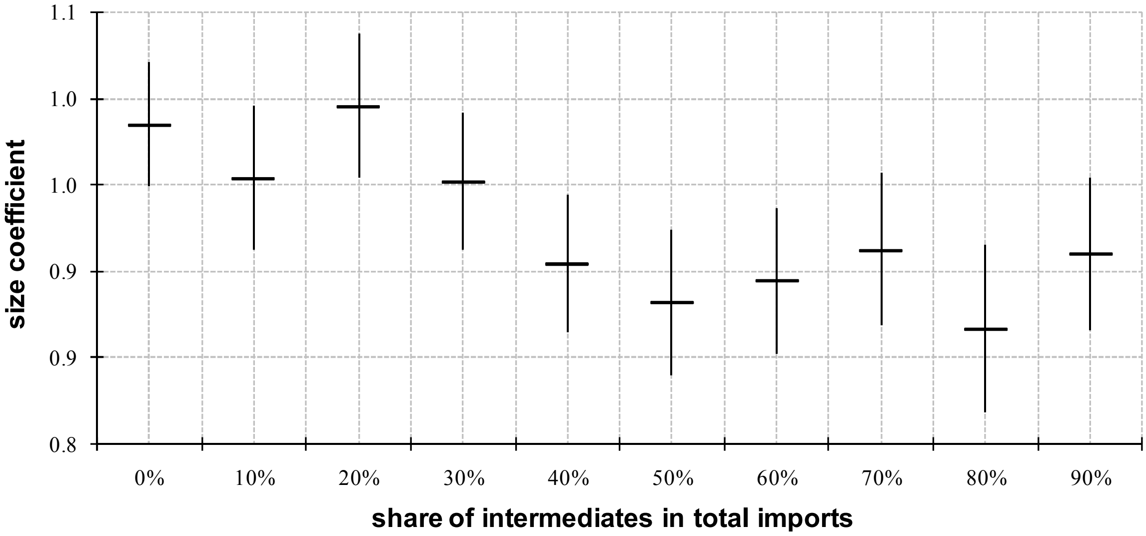
Figure 2
Coefficients by deciles for the standard mass proxy



Note: horizontal bars represent estimated coefficient and vertical bars twice the standard errors.
Source: Authors' estimations.

5. A SEARCH FOR MASS PROXIES WHEN INTERMEDIATES ARE IMPORTANT

The previous section provides clear evidence that the standard gravity equation is 'broken' when it comes to bilateral flows where intermediates trade is important. The theory suggests that the perfect solution would require data on total costs to construct the demand shifter for intermediates imports. If the economy is reasonably competitive, gross sales would be a good proxy for the total costs. Unfortunately, such data are not available for a wide range of nations especially the developing nations where production fragmentation is so important. On the mass variable for the origin nation, theory suggests that we use gross output rather than value added. Again such data are not widely available.

This section presents the results of our search for a pragmatic 'repair' which relies only on data that is available for a wide range of nations. The basic thrust is to use the theory in Section 2 to develop some proxies for economic mass variables that better reflect the fact that the demand for intermediates depends upon gross output, not value-added.

5.1. Fixes for economic mass proxies

We start with the destination nation's mass variable. In Section 2 we showed that a bilateral flow of total goods is the sum of goods whose demand depends upon the importing nation's GDP (i.e. consumer goods) and goods whose demand depends upon the total costs of the sector buying the relevant intermediates. The theory says that our economic mass measure should be a linear combination of two mass measures, not a log-linear combination (see expressions (9) and (7)).

This suggests a first measure that adds imports of intermediates to GDP. The idea here is to exploit the direct definition of total costs as the cost of primary inputs plus the value of intermediate inputs. For any given local firm, some of the intermediates it purchases will be from local suppliers, but summing across all sectors and firms within a single nation, such intermediates will cancel out leaving only payments to local factors of production and imports of intermediates. Our first pragmatic fix therefore is to measure the destination nation's demand shifter by:

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$$E_d \equiv Y_d + \sum_{i \neq o} V_{d,i}^{interm} \tag{11}$$

where V^{interm} is the value of bilateral imports of intermediates. If we summed across all partners, this measure would include part of the bilateral flow to be explained (namely intermediates from nation-o to nation-d). To avoid putting the trade flow to be explained on both sides of the equation, we build the measure for each pair in a way that excludes the pair's bilateral trade.

For the economic mass variable size pertinent to the origin nation, we are trying to capture gross output that must be sold. The proposed measure is a straightforward application of the theory; it uses the origin nation's value-added in manufacturing and its purchases of intermediate inputs from all sources except from itself (due to a lack of data).

$$C_o \equiv AV_o^{manuf} + \sum_{i \neq o} V_{i,o}^{interm} \tag{12}$$

Note that our specification of the gravity equation uses the exports from nation-o to nation-d, so the second term in this does not include the bilateral flow to be explained. The second term involves nation-o's imports from all nations.

5.2. Empirical results

To test whether these proposed proxies work better than GDP, we run regressions like those reported in Table 4 but with the new proxies for economic mass replacing the standard proxy (i.e. GDP). The results are shown in Table 6.

The results in Table 6 – compared with those in Table 4 – suggest that our proxies work better than GDP. The key piece of evidence can be seen in column (1). This includes the ratio of intermediates in total bilateral trade both on its own and interacted with the mass variable. The lack of significance of the ratio in either role suggests that our new proxy is doing a better job than GDP did in picking up demand and supply of intermediates.

Interestingly, the column (2) regression, which allows an interaction between distances on the ratio of intermediates, suggests that the distance coefficient may also be mis-specified. When the ratio is interacted with distance, the distance estimate falls somewhat on average but especially for trade flows where parts and components are especially important (i.e. the ratio is high).

This suggests that distance is more important, not less, for bilateral trade flows dominated by intermediates. The finding may reflect the well-known fact that most production fragmentation arrangements are regional, not global (the parts and components trade is more regionalised that overall trade). This result, however intriguing, does not really stand up to minor changes in the specification. In regression (4), which includes the ratio's interaction with all variables, the distance result fades; indeed only the common language effect seems to be magnified for trade flows marked by particularly high ratios of intermediates.

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Table 6
New mass proxies with share of intermediate, all nations, 2000–2007

Variables	(1)	(2)	(3)	(4)
M_d^{interm}/M_d	1.180	2.644**	2.044**	1.907*
	(1.020)	(1.142)	(0.988)	(1.143)
Ln (E_dC_o/Ω_oP_d)	0.898***	0.889***	0.945***	0.932***
	(0.012)	(0.0116)	(0.012)	(0.012)
* M_d^{interm}/M_d	-0.0322	-0.0132	-0.0289	-0.0247
	(0.020)	(0.020)	(0.020)	(0.020)
ln (Distance$_)$	-1.080***	-0.929***	-0.908***	-0.838***
	(0.018)	(0.038)	(0.019)	(0.038)
* M_d^{interm}/M_d		-0.279***		-0.131*
		(0.065)		(0.067)
Contig$_{od}$			1.441***	1.211***
			(0.092)	(0.224)
*M_d^{interm}/M_d				0.356
				(0.354)
Common language			1.251***	1.047***
			(0.047)	(0.088)
* M_d^{interm}/M_d				0.385***
				(0.143)
Constant	-20.05***	-20.87***	-24.17***	-24.08***
	(0.623)	(0.687)	(0.610)	(0.685)
Observations	87258	87258	87258	87258
R-squared	0.607	0.607	0.631	0.631

Note: Robust standard errors are reported in parenthesis: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$; Pair effects, standard errors clustered by pair; M_d^{interm}/M_d is the share of intermediate imports by a country d over its total imports. New mass variables defined in the text.

Importantly, we note that in all specifications, the ratio's interaction term on the economic mass is always insignificant. This suggests that our new mass proxies are doing a better job of picking up the true supply and demand variables including intermediates.

For symmetry, and to check for non-linear interaction terms, we use our new mass proxies in a regression akin to Table 5. The idea is to use ratio decile dummies instead of the ratio itself in order to allow the interactions to vary non-linearly for bilateral flows marked by different degrees of intermediates trade. The results are shown in Table 7.

To interpret our findings, recall that the significance of the upper-tier decile interaction terms was taken as evidence that GDP was not working well for trade flows marked by much trade in intermediates. Thus the results in Table 7 suggest that our new proxy is working better than GDP.

Specifically, the base-effect for our economic mass variable and the distance coefficients are estimated at very reasonable point estimates (0.88 and -1.1 respectively). Critically, only one of the decile interaction terms is significant, and it is positive, not negative as the theory would suggest. Two other interaction terms are borderline significant and negative, the ones for the sixths and tenth deciles.

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Table 7
New mass proxies with intermediate deciles, all nations, 2000–2007

	Ln (E_dC_o/Ω_oP_d)	**ln (Distance)**	**Constant**
Base effect	0.877***	-1.051***	-19.29***
	(0.022)	(0.018)	(1.074)
Base effect * d2	0.0402		
	(0.024)		
Base effect * d3	0.0365***		
	(0.025)		
Base effect * d4	0.0294		
	(0.024)		
Base effect * d5	-0.0256		
	(0.024)		
Base effect * d6	-0.0531**		
	(0.025)		
Base effect * d7	-0.0390		
	(0.025)		
Base effect * d8	-0.0306		
	(0.026)		
Base effect * d9	-0.0652**		
	(0.028)		
Base effect * d10	0.0102		
	(0.027)		
Observations	87251		
R-squared	0.609		

Notes: See notes to Table 5.

6. WHY DO INCORRECTLY SPECIFIED MASS VARIABLES MATTER?

A large number of gravity studies focus on variable that vary across country pairs – say free-trade agreements, cultural ties, or immigrant networks. The most recent of these studies employ estimators that control for the mass variables with fixed effects.[5] Such studies do not suffer from mass-variable mis-specification and so are unaffected by our critique.

There are however as mentioned in the introduction, a number of recent studies – especially concerning the 'distance puzzle' that do proxy for the production and demand variables with GDP. It is these studies that our work speaks to.[6]

However, since most of these studies are concerned with a broad set of nations and commodities, the mis-specification of the mass variable probably has a minor impact on the

5 These econometric techniques were introduced by Anderson and Van Wincoop (2003), and Combes, Lafourcade and Mayer (2005), Feenstra (2003), Harrigan (1996), and Head and Mayer (2000).

6 Rauch (1999), Brun et al. (2005), Berthelon and Freund (2008), Jacks et al. (2008), and Anderson and Van Wincoop (2003).

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results – as the findings of Bergstrand and Egger (2010) showed and we confirmed with our Table 1 results. More worrying, however, is the use by authors that focus on trade in parts and components.[7] These papers use the consumer-good version of the gravity model and thus mis-specify the mass variable.

Once the equation is mis-specified – in particular the standard economic mass proxies are not correctly reflecting the supply and demand constraints – we are in the realm of omitted variable biases. The first task is to explore the nature of the biases that would arise from this mis-specification. To simplify, assume away GDPs and distance and focus on a pair-wise policy variable, say, nation-d's tariffs on imports from nation-o; we denote this as T_{od}. The estimated gravity equation will have the following structure:

$$\ln V_{odt} = \text{constant} + a_5 \ln T_{odt} + \varepsilon_{odt} \tag{13}$$

where the error is assumed to be iid.

Because intermediates supply is measured by total costs rather than GDP, and the supply of intermediates that must be sold depends upon gross output rather than value added. This means that the true model includes an additional term. That is:

$$\ln V_{odt} = a_0 + a_5 \ln T_{odt} + a_6 \ln Z_{odt} + \varepsilon_{odt} \tag{14}$$

where Z_{odt} is the difference between the GDP-based mass variables and the true mass variables as specified in (7). We can write Z_{odt} as a function of T_{odt} in an auxiliary regression:

$$\ln Z_{odt} = b_0 + b_1 \ln T_{odt} + u_{odt} \tag{15}$$

where u is assumed to be iid. Using this notation for the coefficients of the auxiliary regression, we can see that in estimating (13), we are actually estimating:

$$\ln V_{odt} = a_0 + a_5 \ln T_{odt} + a_6 \ln Z_{odt} + \varepsilon_{odt} \tag{16}$$

What this tells us is that the coefficient on the policy variable of interest will almost certainly be biased. The point is that the only way it is not biased is if there is no correlation between the mis-specification of the economic mass variables and the policy variable.

What sort of correlation should we expect? Recall that the mis-measurement of the economic mass variable all goes back to the importance of trade in intermediate goods. Since almost all bilateral variables of interest are things that affect bilateral trade flows, it seems extremely likely that the variable of interest will also affect the flow of intermediates. As long as it does, then we know that the mis-specification of the mass variable will also lead to a bias in the pair-wise variables.[8]

For example, let us suppose that tariffs discourage trade overall, but they especially discourage intermediates trade (for the usual effective rate of protection reasons, i.e. the tariff is paid on the gross trade value but its incidence falls on the value-added only). In this case, we should expect low tariffs to encourage two things, an overall increase in trade and an increase in the ratio of intermediates. In this case, the bias in the mis-specified gravity equation is likely to be negative, since the policy variable is negatively correlated with the omitted variable. Furthermore, the mis-specification also affects the standard errors, which would result in a biased inference (Wooldridge, 2003, ch.4).

[7] Athukorala and Yamashita (2006), Kimura et al. (2007), Yokota (2008), and Ando and Kimura (2009).

[8] As noted above, the modern techniques for controlling for mass with time-varying country-specific dummies eliminates such biases since they correctly control for the role of intermediates.

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7. CONCLUDING REMARKS

In this paper we present empirical evidence that the standard gravity model performs poorly by some measures when it is applied to bilateral flows where the parts and components trade is important. The paper also provides a simple theoretical foundation for a modified gravity equation that is suited to explaining trade where international supply chains are important. Finally we suggest ways in which the theoretical model can be implemented empirically.

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APPENDIX

Classification for intermediate and final goods

	BEC categories
Intermediate goods:	111 – Primary food and beverages, mainly for industry
	121 – Processed food and beverages, mainly for industry
	21 – Primary industrial supplies not elsewhere specified
	22 – Processed industrial supplies not elsewhere specified
	32 – Processed fuels and lubricants
	42 – Parts and accessories of capital goods (except transport equipment)
	53 – Parts and accessories of transport equipment
Consumption goods:	112 – Primary food and beverages, mainly for household consumption
	122 – Processed food and beverages, mainly for industry
	51 – Passenger motor cars
	6 – Consumer goods not elsewhere specified
Other:	31 – Primary fuels and lubricants
	41 – Capital goods, excluding parts and components
	51 – Other transport equipment
	7 – Other

Source: Comtrade's Broad Economic Categories; for details see http://unstats.un.org/unsd/tradekb/Knowledgebase/Intermediate-Goods-in-Trade-Statistics.

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Household Money Holdings in the Euro Area: an explorative investigation*

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ABSTRACT

In this paper we analyse household holdings of the broad monetary aggregate M3 in the euro area from 1991 until 2009. Households are the largest money-holding sector in the euro area. We develop four models, two in nominal, two in real terms, with satisfactory economic and statistical properties. The main determinants are a transactions variable, wealth considerations, opportunity costs and uncertainty. In particular housing wealth is found to play an important role. The models are robust to different estimation strategies, samples considered and a multitude of misspecification tests. According to our analysis, it is quite apparent that in equilibrium, households jointly determine consumption and broad money holdings which are both influenced by wealth as well as interest rates. The importance of household money holdings for consumption expenditures may cast doubt on a purely passive role for money.

Keywords: money demand, cointegrated VARs, households

JEL Classification: E41, C32, D12

1. INTRODUCTION

Understanding the demand for money is an important element of a detailed monetary analysis, which aims to extract, in real time, signals from monetary developments that are relevant for the assessment of risks to price stability over the medium to longer term. These longer-term price developments are determined by aggregate money holdings of all sectors. Looking at individual

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money holding sectors, however, may allow the formulation of more consistent and richer explanations of the driving forces for the demand for money as the relative importance of the main motives for holding money - its use as a medium of exchange or as a store of value - varies across sectors. Indeed, heterogeneity in money-holding behaviour goes beyond the sector level to the individual money holder, but harmonised data for a significant sample length is only available at the sectoral level.[2] A sectoral analysis of money demand can, by improving the understanding of the individual components, contribute to a better understanding of the covariation of aggregate money with its determinants.

In general, differences in money demand behaviour may result from two factors:

1. The constraints surrounding the money-holding decision process can vary. This may lead to different elasticities of money demand with respect to the same determinants for individual sectors.
2. The determinants of money demand may differ across sectors, such as alternative investment opportunities and thus different opportunity costs of holding money, or different scale variables.

Consequently, two different modelling strategies need to be considered in the context of sectoral money demand. The first is to estimate money demand using a common set of macroeconomic determinants (see von Landesberger, 2007). This approach allows for a comparison of behaviour across sectors and with aggregate money demand. The alternative modelling approach is oriented toward finding a refined specification for every sector, thus trying to identify the determinants best capable of explaining sectoral money holdings. This is the aim of the present paper which has not yet been done for euro-area data. The understanding of household money holdings is important for several reasons: households are the largest money-holding sector accounting for approximately two-thirds of euro-area M3. They usually hold a large proportion of their money holdings as transactions balances, using these balances mainly as a buffer, while slowly adjusting their portfolio composition. In addition, households' financial decisions are likely to have significant impact on real macroeconomic activity, rendering the interaction between households' money balances and consumption important. The dynamic of household M3 holdings are also found to be informative for price developments in the euro area, giving their explanation a particular relevance for monetary analysis (see European Central Bank, 2006 p. 18).

The paper is structured as follows. The next section provides a review of the literature on household money demand. In part 3, the data and the modelling approach used to estimate the money demand systems are discussed. We also present and discuss the results of four different models of households' M3 demand. Section 4 illustrates the use of these models to understand recent money growth. As a robustness check, section 5 also presents results of a single-equation modelling approach. The last section summarises the findings and provides some implications for monetary analysis.

2. RELATED STUDIES

The following section provides a structured overview of the methods commonly employed in the literature on sectoral money demand and of the main findings. In order to get a better understanding of the results, we distinguish between macroeconomic (time-series) and microeconomic (cross-sectional) studies.

[2] See Martinez-Carrascal and von Landesberger (2010) for a comparison of the behaviour of money demand at the sectoral level and at the micro-economic level for euro-area non-financial corporations.

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2.1. The evidence on household money demand from macroeconomic studies

For the US, the first empirical analysis of the household demand for money was undertaken by Goldfeld (1973). In this study, the demand for M1 is explained by different measures of transactions (GNP and consumption expenditure), controlling for the change in net worth and using the spread between commercial paper and deposit interest rates as opportunity costs. Goldfeld finds that money holdings by households are quite well explained by these variables and have reasonable parameter estimates. Since the publication of Goldfeld (1973), a number of studies have attempted to explain household money demand. In general, these studies have analysed the demand for money by households from a time series perspective using cointegration methods – either based on single equations (Butkiewicz/McConnell, 1995) or based on systems of equations (e.g. Jain and Moon, 1994, Thomas, 1997, Chrystal/Mizen, 2001).

The main scale variable of money demand considered includes real consumer expenditure (Jain and Moon, 1994, Read, 1996, Calza and Zaghini, 2010), measures of real (permanent) disposable income (Butkiewicz and McConnell, 1995, Laumas, 1979), real net labour income (Chrystal and Mizen, 2001) and real GDP (Petursson, 2000, Feiss and MacDonald, 2001). In addition, both real gross personal sector wealth (Thomas, 1997, Read, 1996) and real net total wealth (Chrystal and Mizen, 2001) are intended to capture an additional element of scale.

A variety of interest rate specifications have been tried. These range from simple formulations such as including only the long-term nominal treasury yield (Jain and Moon, 1994), the short-term commercial paper rate (Laumas, 1979), or foreign interest rates (Krupkina and Ponomarenko, 2013). Semi-log and double-log specifications are used (Butkiewicz and McConnell, 1995, Calza and Zaghini, 2010).[3] More complex approaches include the spread between the 3 month t-bill rate and the own rate of money (Thomas, 1997, Petursson, 2000) or between the yield on public bonds and the own rate (Read, 1996). Chrystal and Mizen (2001) even include two interest rate terms in their model, the rate on savings deposits minus a money-market rate and the spread between the rate on consumer credit and the base rate. An additional variable repeatedly included in models for the UK is the rate of inflation, reflecting either the return on real alternatives to money or helping to test for price homogeneity (Thomas, 1997, Chrystal and Mizen, 2001, Feiss and MacDonald, 2001).

The main findings are that household real balances are cointegrated with measures of income and interest rates. Several studies emphasise, both for narrow and for broad monetary aggregates, a transactions-based explanation of money demand (Jain and Moon, 1994), captured by a strong interaction between household money holdings and consumption (Thomas, 1997). Broadening the analytical framework to include households demand for loans, Chrystal and Mizen (2001) find that consumption, money holdings and credit interact both in the determination of the long-run equilibrium and in their short-run adjustment. Read (1996) provides evidence for Germany that households' money holdings tend to be determined by longer-term considerations, whereas the corporate sector is far more responsive to short-term influences.

2.2. Microeconomic evidence at the household level

While the focus of this paper is a time series perspective, evidence brought forward in cross-sectional studies could potentially contribute valuable further insights in the specification of the models. The monetary data examined in these studies is generally taken from household surveys. A first study was conducted by Garver and Radecki (1987) on a cross-sectional sample of US data. They investigate the holdings by households of a narrow measure of money consisting of currency holdings plus total checking accounts. The scale variable considered is total household

[3] Calza and Zaghini (2010) find that the welfare costs of inflation differ for households and firms and crucially depend on the double-log versus semi-log specification.

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annual income, while the opportunity costs of holding money are measured by the average money-market deposit rate minus the rate of interest earned on checking accounts. A number of dummy variables are included to account for different types of checking accounts. The study emphasises the transactions motive for holding money and support the use of the macroeconomic approach to the demand for narrow money. Attanasio et al. (1998) also investigate households' holdings of real cash balances using non-durable consumption as scale variable and an interest rate as opportunity costs. The interest rate and expenditure elasticities found for the demand for cash are close to the theoretical values implied by standard inventory models. With data for Japan, Fujiki and Hsiao (2008) constitute an exception by examining the issues of unobserved heterogeneity among cross-sectional units and stability of an aggregate function for broad money. The estimated income elasticity for Japanese household M3 is around 0.68 and the five year bond interest rate elasticity is about -0.12. Anderson and Collins (1997) investigate M2 growth in the United States from 1990 to 1993 using a model of household demand for liquid wealth. The authors find that the own-price elasticity of money demand rose substantially during this period and report sizeable cross-price elasticities of money with respect to other liquid financial assets, notably with mutual funds. They also suggest that households may respond more rapidly to changes in market interest rates than is often assumed. Tin (2008) examines the precautionary demand for transactions balances. The monetary measure considered is non-interest-earning checking accounts by households in the US. The study indicates that income volatility is a significant determinant of money holdings as predicted by the inventory theory of money demand. The relative magnitudes of the elasticities of income and income volatility suggest that the strength of the relationship between the precautionary motive and money demand is much weaker than the strength of the relationship between the transactions motive and money demand. Unfortunately, cross-sectional studies have not yet investigated holdings of components of broad monetary aggregates.

3. THE EMPIRICAL APPROACH

In what follows, we try to model euro area M3 holdings by households using a broad set of explanatory variables. In line with the literature reported above, the aim is to estimate a money demand relationship using system cointegration techniques.

3.1. The framework and the data

Monetary theory suggests different determinants for the holding of broad money, which like for other financial assets, is part of a portfolio allocation decision (see Friedman, 1956, Tobin, 1969). At least some of the assets included in broad money in addition provide liquidity services to their holder. A general formulation of the determinants can be stated:

$$m = \beta_1 p + \beta_2 y + \beta_3 w + \beta_4 i^{alt} + \beta_5 i^{own} + \beta_6 \sigma$$

whereby m denotes the stock of money, p the price level, y and w the level of transactions and wealth, respectively, i^{alt} and i^{own} the returns for investments outside M3 and in monetary assets included in M3 and σ represents variables capturing different aspects of uncertainty, be they economic, financial or geopolitical. The β_i's ($i = 1,\ldots,6$) denote the parameters capturing the effect of the respective determinants on money holding.

Three key economic features have to be fulfilled by empirical estimates in order to identify a money demand function:

1. β_2 and β_3 must be positive,
2. β_4 must be negative and β_5 positive,

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3. discrepancies between actual money and equilibrium holdings lead to an adjustment in money growth.

For the purpose of our analysis, *m* is broad money holdings of households. The sector also comprises non-profit institutions serving households.[4] M3 data is taken from the official ECB database for the period since 1999.[5] The criticism is sometimes made that sectoral money demand studies for the US, which are based on flow-of-funds data, might be affected by the fact the household money holdings are a residual position in the data. In the euro area, over the sample considered, this is not the case, as between 81% and 88% of M3 data, namely all deposits (including repurchase agreements) held by the household sector were directly reported by MFIs.

Two measures of the price level are considered potentially relevant for households to calculate real balances: the private consumption deflator and the harmonised index of consumer prices. The scale of households' transactions settled using money may be captured by a variety of variables. Following the literature, the variables considered are the level of consumption expenditures, disposable income, a measure of household expenditures consisting of consumption plus investment into housing as well as measures of household wealth. Boone et al. (2004), Greiber and Setzer (2007), Beyer (2009) and de Bondt (2009) find a significant role for wealth in euro area money demand, with the latter three studies emphasising the role of housing wealth. A number of wealth measures are therefore considered, specifically gross total household wealth and its components gross financial household wealth and gross housing wealth, as well as net total household wealth. In addition, a measure of longer-term housing wealth using the trend in house prices is used. The reason underlying such a calculation is that households may not perceive themselves to become more or less wealthy with high frequency movements in the prices of their asset holdings, but rather take a medium term-view of asset prices.[6]

In order to model the opportunity costs of holding money, a wide range of alternative returns and interest rates are initially selected: these include the long-term interest rate on bank lending to households for house purchase, the yield on long-term government bonds, the yield on corporate bonds and a short-term money market interest rate. While the last three interest rates can be seen as fairly common choices, the consideration of the bank lending rate as an alternative investment opportunity for households rests on the observation that in the presence of intermediation costs between borrowing and lending from a bank, a reduction in borrowing generally offers households a better return than holding money. Thus, the use of a bank lending rate draws on the notion that the household sector holds money as a buffer stock which will be reduced as the financing cost of households increases.

What is left are proxies for risky investments: the dividend yield as well as the earnings yield of euro area non-financial corporations are considered to take developments on stock markets into account. Friedman (1988) outlines the interactions between money holdings and the stock market. In addition to the realised earnings per share, following the approach proposed by Chordia and Shivakumar (2002) as well as Stern and Stern (2008), expected earnings per share are estimated based on a regression relating the earnings on equity on the recent dividend yield, the real short-term interest rate, the slope of the yield curve and the spread between corporate and government bonds. Moreover, the simple price/earnings ratio is also included in the data set. It is determined by expectations about discount rates and about earnings growth, with the former mainly influencing the evolution in the long-run (see Fama and French, 2002, Campbell and Shiller, 1998). It can therefore be considered as a proxy for the discount rate applied to investments in risky assets. The spread between corporate and government bonds can also be viewed as a proxy for risk. The

[4] The level of money stock is the notional stock adjusted for seasonal effects with Tramo-Seats. The data is extended backwards before 1999 Q1 assuming an unchanged sectoral share in money market funds, currency in circulation and debt securities holdings at the levels of 1999 Q1. These instruments represent only a small share of household M3 holdings in 1999.

[5] The overall approach to the construction of the series is described in ECB (2006b).

[6] The trend in house prices is derived using an approach common to the analysis of the link between money and asset prices (Detken and Smets, 2004, Adalid and Detken, 2007). The trend is estimated using a very slow adjusting HP-Filter ($\lambda = 100{,}000$).

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return on monetary assets is captured by the own rate of households' M3 holdings, calculated as a weighted average of the remuneration of the instruments included in M3.

Interest rates can enter the money demand relationship in two functional forms: first, the semi-log specification, which is the most popular in money demand studies (see, e. g., Ericsson, 1998). It estimates semi-elasticities and implies the same response of money holdings to each percentage point reduction in nominal interest rates. Second, the double log form proposed, inter alia, by Lucas (2000). It entails that a percentage point reduction in nominal interest rates has a proportionally greater impact upon money holdings the lower the level of interest rates, i.e., the semi-elasticities vary with the level of interest rates. For higher levels of interest rates the two functional forms lead to similar results. The non-linear impact at low levels of interest rates can be motivated by the prevalence of fixed costs into alternative investment opportunities and that households who hold only cash do not incur this cost. A logarithmic money demand function may also be rationalized within a stylised general equilibrium model with money (Chadha et al., 1998, Stracca, 2001).

Finally, measures of uncertainty that proxy households' economic and financial confidence are also included. The measures considered are the EU Commission's index of consumer confidence and its subcomponents (e.g. employment expectations) as well as the actual rate of unemployment, which was found relevant by de Bondt (2009). Furthermore, financial market uncertainty in the form of stock and bond market volatilities (Carstensen, 2006) and of the uncertainty factors estimated by Greiber and Lemke (2005) is taken into account. The latter derive composite series for uncertainty using an unobserved components model. One of their indicator variables is mainly based on financial market data, such as medium-term returns, loss and volatility measures while the other factor is more heavily geared toward business and consumer sentiment. Both the individual economic variables as well as the aggregate factors are intended to capture the economic forces impacting on the household's decision to hold money for precautionary reasons.

The set of explanatory variables presented above and shown in Chart A in annex 1 allow the specification of a whole battery of equations. A number of alternative specifications for household M3 holdings are tested and a selection of the most promising specifications is presented in more detail below. The equations are chosen to get economically plausible specifications and statistically sound estimation results. More specifically, the equations considered are:

Model 1-n: $$m = f\left(\overset{+}{pc},\ \overset{+}{rc},\ \overset{+}{rthw},\ blr \overset{-}{-} own,\ \overset{+}{GL1},\ \overset{?}{UN^e}\right)$$

Model 2-n: $$m = f\left(\overset{+}{pc},\ \overset{+}{rc},\ \overset{+}{rtw},\ blr \overset{-}{-} own,\ p \overset{-}{-} e\right)$$

Model 3-r: $$m - pc = f\left(\overset{+}{rdi},\ \overset{-}{dpc},\ IRL \overset{-}{-} IRS,\ p \overset{-}{-} e\right)$$

Model 4-r: $$m - pc = f\left(\overset{+}{rc},\ \overset{-}{dpc},\ IRL \overset{-}{-} OWN,\ p \overset{-}{-} e,\ \overset{?}{C^e}\right)$$

where variables written in lower case letters enter the VAR systems in logarithms. The sign above the variables indicates the theoretical expected impact.

Model 1-n explains *nominal* household M3 holdings using the private consumption deflator *pc* and two scale variables, real private consumption *rc* and a measure of the trend in housing wealth deflated with the private consumption deflator *rthw*. The spread between the bank lending rate for house purchases *blr* and the own rate on households' M3 holdings *own* (both in logs) enters the money demand model as the measure of opportunity costs. In order to model precautionary motives of the demand for money, the uncertainty measure developed in Greiber and Lemke *GL1*

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related to capital market forces enters the model as a measure of uncertainty. Finally, expectations with regard to unemployment over the coming twelve months UN^e from the survey of the EU Commission are included in the VAR system. A deteriorating employment situation may, on the one hand, induce households to hold greater money balances to meet unforeseen expenditures. On the other hand, the expected deteriorating economic environment and increasing uncertainty may reduce the attractiveness of nominal assets and induce the purchase of more real assets. Therefore, the total impact on the demand for money is ambiguous (see Atta-Mensah, 2004b).

Model 2-n draws on a similar set of variables as Model 1-n, but includes total household wealth deflated with the private consumption deflator as the relevant wealth measure *rtw*. Precautionary money holdings are captured by the price earnings ratio on euro area equity $p - e$ which may be considered as a measure of risk on financial markets.

Model 3-r explains *real* household M3 balances with only one scale variable – real disposable income *rdi*, but includes two measures of opportunity costs, the change in the consumption expenditure deflator *dpc* and the term spread *IRL – IRS*. As in Models 2-n, the price-earnings ratio $p - e$ is also included.

Model 4-r builds on the previous model, but substitutes *rdi* with real consumption expenditure *rc*, *IRL – IRS* with the spread between the long-term nominal bond yield and the own rate of household M3 holdings *IRL – OWN*. In addition, expectations of economic prospects and thus future consumption are taken into account. In order to capture this forward-looking element, expectations with regard to the strength of economic activity from the EU Commission consumer confidence surveys C^e are included.

3.2. Overview of the modeling outcomes

The empirical analysis is conducted on seasonally adjusted quarterly data over the sample period 1991 Q1 to 2009 Q3.[7] The estimations are performed over the shorter sample 1991 Q1 to 2008 Q3 in order to avoid any contamination of the results from the financial market crisis following the default of Lehman Brothers in September 2008, with the last four observations analysed in Section 4.

To determine the order of integration of the time series, ADF and KPSS tests are carried out (see Table 9 in annex 1). The two tests – together with conceptual considerations for some of the borderline cases on the boundedness of the variance - support the view that most series in levels, except the spreads, are I(1). An additional test for stationarity of the variables within the cointegrated VAR supports this decision (see Table 11 in annex 1). That said, it should be recognised that some variables may still exhibit quite persistent fluctuations in first differences. Difference stationarity of money and prices may be considered slightly at odds with a part of the recent empirical literature on money demand that finds these variables to be I(2).[8] Given its prominence in the empirical money-demand literature, this possibility is entertained in the modelling approach applied below.

Within our cointegrated VAR approach the first step consists in estimating an unrestricted VAR system comprising an endogenous variables vector y_t and exogenous (non-modelled) I(0) variables vector x_t:

[7] Davidson and MacKinnon (1993, p. 714) prove that unit root test statistics are biased against rejecting the null hypothesis when working with seasonally adjusted data. As nearly all our variables are clearly *I(1)* (see annex 1) this reduces the severity of this problem. Furthermore, Ericsson et al. (1994) show theoretically and empirically within the Johansen framework that the number of cointegrating vectors and the cointegrating vectors themselves are invariant to the use of seasonally adjusted or unadjusted data.

[8] See Juselius (2006) and for instance Feiss and MacDonald (2001).

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$$y_t = \sum_{i-1}^{p} \Pi_i y_{t-i} + \Psi_0 x_t + \Phi D_t + \varepsilon_t \tag{1}$$

The errors ε_t are assumed to be $NI\text{\textasciitilde}(0,\Omega)$. Π_i and F are matrices containing the parameters of the model. D_t is a vector of deterministic variables, potentially comprising constant terms μ_0 or deterministic trends. Given the quarterly data used, the maximum lag length p is set equal to four in order to determine the appropriate number of lags for each model. The Akaike information criterion (AIC) is used to select the lag length for conducting the remainder of the analysis and the outcome is cross-checked with Likelihood Ratio tests (see Table 10 in annex 1). The AIC tends to favour the inclusion of more lagged terms than for example the Schwartz information criterion.[9] Overestimation of the order of the VAR is much less serious than underestimating it, as shown for example by Kilian (2001). In the models presented below, the lag length retained ranges between two and three in levels.

Table 1 presents the outcome of standard specification tests of the respective VAR systems. The null of no autocorrelation in the residuals cannot be rejected in any of the systems at conventional significance levels. In a similar vein, tests for ARCH effects in the residuals are also not significant. By contrast, non-normality of the residuals is detected for two models owing to the presence of outliers.[10]

Table 1
Residual properties for the VAR systems

	Test statistic	p-value		Test statistic	p-value
Model 1-n					
LM-AR(1)	F(36,54) = 0.69	0.95	Multivariate ARCH	F(441,61) = 1.11	0.32
LM-AR(4)	F(36,51) = 0.99	0.50	Normality	F(12,59) = 1.67	0.14
Model 2-n					
LM-AR(1)	F(36,48) = 1.16	0.31	Multivariate ARCH	F(441,55) = 1.13	0.31
LM-AR(4)	F(36,44) = 0.59	0.95	Normality	F(12,53) = 2.27	0.02
Model 3-r					
LM-AR(1)	F(25,57) = 1.53	0.09	Multivariate ARCH	F(225,50) = 1.32	0.09
LM-AR(4)	F(25,54) = 0.88	0.63	Normality	F(10,61) = 0.79	0.64
Model 4-r					
LM-AR(1)	F(36,53) = 0.96	0.55	Multivariate ARCH	F(441,60) = 1.07	0.39
LM-AR(4)	F(36,50) = 0.65	0.91	Normality	F(12,58) = 2.58	0.01

Note: p-values derived from comparison with respective asymptotic distribution.

In a second step, we reformulate the VAR system into a VECM and test for the rank of the matrix Π_1 using the trace test (see Johansen, 1996):

$$\Delta y_t = \Pi_1 y_{t-1} + \mu_0 + \sum_{i=1}^{l-1} \Gamma_i \Delta y_{t-i} + \Psi_0 x_t + \varepsilon_t \tag{2}$$

where l indicates the lag length determined in the previous step. The trace tests were conducted assuming the presence of a linear deterministic trend in the time series and a non-zero intercept

[9] Lütkepohl and Saikonnen (1997, p. 16) find that "In most cases AIC and HQ have a slight advantage over the very parsimonious SC criterion".

[10] While normality of residuals is part of the theoretical assumptions of the distribution of residuals, the violation of normality may not be a severe deficiency as the evaluation of the trace test will be supported by bootstrapping results.

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m_0 in the cointegration relationship.[11] Table 2 reports the trace test statistics for different rank assumptions as well as the p-values obtained from comparing this test statistic with the critical values derived by MacKinnon et al. (1999).[12] All models reject the rank 0 at the 5% significance level, with model 1 also rejecting rank 1 and 2. However, given the presence of exogenous I(0) regressors in one of the models (in model 1-n unemployment expectations) and the small sample size, caution in assessing the number of long-run relationships possibly present in the data using this metric seems reasonable. Therefore, more informative parametrically bootstrapped p-values generated from 1,000 replications are undertaken.[13] While the theory on bootstrapping in a non-stationary framework, such as the cointegrated VAR, is still undiscovered territory, the usual theoretical properties from models with stationary variables seem to apply in this setting as well (Juselius, 2006, p. 157, Swensen, 2006).[14] At the 10% significance level, all model specifications indicate one cointegration relationship.

System cointegration tests are well-known to have low power. This gives reason to believe that such tests have a tendency to favour the choice of too few long-run relations. Juselius (2006) suggests the use of as much additional information as possible in the rank determination. We follow this lead and additionally:

1. examine whether the t-value on the load factor of an additional cointegration vector is less than 2.6;
2. analyse recursively the trace statistic and the cointegration relations;
3. check the economic interpretability of the results.
4. While the first and third approaches require the specification of the cointegrated VAR systems, the second approach can be generated on the basis of the unrestricted VAR model.

Table 2
Trace test results

Model		Rank					
		0	1	2	3	4	5
1-n	test statistic	133.56	83.98	46.95	13.22	5.71	0.83
	p-value	0.000	0.001	0.020	0.671	0.475	0.362
	bootstrapped p-value	0.022	0.213	0.473	0.961	0.792	0.658
2-n	test statistic	119.16	53.99	31.21	15.61	7.11	1.90
	p-value*	0.000	0.462	0.655	0.738	0.565	0.17
	bootstrapped p-value	0.074	0.798	0.785	0.957	0.854	0.292
3-r	Test	75.77	37.75	14.07	6. 52	1.863	-
	p-value*	0.016	0.313	0.837	0.633	0.172	-
	bootstrapped p-value	0.012	0.263	0.789	0.712	0.215	-
4-r	Test	98.93	58.13	31.71	15.41	4.42	0.32
	p-value*	0.029	0.297	0.628	0.752	0.866	0.573
	bootstrapped p-value	0.052	0.336	0.601	0.768	0.867	0.536

* Barlett corrected trace statistic.

[11] The cointegration analysis and the results presented in the remainder of this note are computed with the Structural VAR software which was kindly provided by Anders Warne. See http://www.texlips.net/svar/source.html.

[12] Where no exogenous I(0) regressors are included in the VAR systems, the Bartlett correction of the test statistic is undertaken and compared with the critical value.

[13] The parametric bootstrapping procedure implies drawing new innovations from a multivariate standard normal distribution. These innovations are then transformed into bootstrapped residuals using the estimated covariance matrix from the original estimated residuals. On the basis of the initial values and taking the estimated parameters as given, new data series are constructed and the model re-estimated on the new data set. An alternative would be to adjust the test statistics (see, e. g., Reimers, 1991) or the critical values (see Cheung and Lai, 1993).

[14] In particular, a bootstrapped statistic can be expected to have errors in null rejection probabilities that are of a smaller order of magnitude, as the sample size goes to infinity, than its asymptotic analogue when the asymptotic distribution of the statistic is invariant to the parameters of the model. Almost all statistics that we bootstrap are invariant in this sense. See Park (2005) and Chang et al. (2002) for some recent developments regarding models with unit roots.

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In order to robustify the modelling decision on the basis of the trace test, Chart B in annex 1 shows the recursively estimated trace test statistic for the hypothesis of rank one. For the nominal models, the outcome is more reassuring than for the models specified in real terms. But in any case, the trace statistic is fairly stable and around the level of the critical value for the 5% confidence level. Therefore, in the following, a rank of one is assumed for modelling the VAR systems. This decision is also supported by Chart C in annex 1, which presents recursive estimates of the largest eigenvalue for a given set of parameters of the short-run and deterministic variables. For all four models, the depicted eigenvalue bands do not cross the zero line.

Parameter stability has been an issue of primary concern in the context of money demand estimations. Table 3 presents the outcome of tests on parameter non-constancy under the retained assumption that the P-matrix has rank 1. The Ploberger-Krämer-Kontrus (1989, henceforth PKK) fluctuation test examines the constancy of the parameters capturing the short-run dynamics. The test is conducted for all individual equations of the VAR system, but the table reports only the outcome of the money demand equation. The PKK test is unable to reject the null of parameter constancy which supports the eigenvalue analysis reported above. In addition, Table 3 also shows the results of the Nyblom tests for possible non-constancy of the parameters of the cointegration vector. Again, the stability of the parameters is not rejected.

Table 3
Stability tests – Nyblom & PKK (Rank 1)

Model 1-n	Test statistic	p-value
Nyblom Sup F	87.80	0.16
Nyblom Mean	21.51	0.13
PKK S(9)	0.66	0.96
Model 2-n		
Nyblom Sup F	24.99	0.25
Nyblom Mean	6.06	0.31
PKK S(14)	1.72	0.28
Model 3-r		
Nyblom Sup F	3.75	0.31
Nyblom Mean	0.73	0.58
PKK S(7)	0.56	0.98
Model 4-r		
Nyblom Sup F	28.39	0.54
Nyblom Mean	8.52	0.54
PKK S(8)	0.86	0.76

Note: p-values derived from comparison with respective bootstrapped distributions. PKK is calculated on equation-by-equation basis.

Moreover, tests on the stationarity of the variables included in the VAR are conducted to determine whether the reduced rank of the P-matrix resulted from the inclusion of stationary variables. Table 11 in annex 1 reveals that the reduced rank does not seem to result from a single stationary variable. This again supports our decision to treat the variables as I(1).

In order to get further insights into the relationship between the variables and to help the identification of the cointegrated VAR system, we run joint weak exogeneity tests on the variable set. The tests also help to detect the common driving forces amongst the variables of the system.

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A weakly exogenous variable contributes to the common trend of the other variables in the VAR system. At the same time, shocks to variables that are not weakly exogenous have no permanent effect on any other variable in the system. Table 4 shows the outcome of this analysis. In all four models, the assumption that the cointegration relationship does not affect household M3 balances is clearly rejected (see for further discussion the presentation of the models below).

Following the choice for the rank of the Π_1-matrix in (2), finally α cointegrated VAR system is estimated. This entails the identification and estimation of the vector of load factors a and the cointegration vector β' in (3)

$$\Delta y_t = \alpha\beta' y_{t-1} + \mu_0 + \sum_{i=1}^{l-1} \Gamma_i \Delta y_{t-i} + \Psi_0 x_t + \varepsilon_t \tag{3}$$

Table 4
Tests for weak exogeneity of variables

	Null hypothesis: α in equation k is zero Alternative hypothesis: α in equation k is not zero							
	Model 1-n		**Model 2-n**		**Model 3-r**		**Model 4-r**	
Equation for	F(1,58)	p-value	F(1,52)	p-value	F(1,60)	p-value	F(1,59)	p-value
m	26.85	0.00	13.35	0.00	-	-	-	-
$m - pc$	-	-	-	-	5.16	0.03	15.33	0.00
pc	0.60	0.44	13.03	0.00	-	-	-	-
Rc	10.49	0.00	16.72	0.00	-	-	0.17	0.68
Rdi	-	-	-	-	20.25	0.00	-	-
Rtw	-	-	5.19	0.027	-	-	-	-
$Rthw$	1.81	0.18	-	-	-	-	-	-
Dpc	-	-	-	-	5.04	0.03	0.17	0.68
$blr - own$	0.41	0.52	7.66	0.01	-	-	-	-
$IRL - IRS$	-	-	-	-	1.51	0.22	-	-
$IRL - OWN$	-	-	-	-	-	-	1.39	0.24
$GL1$	0.42	0.52	-	-	-	-	-	-
$p - e$	-	-	0.12	0.73	4.15	0.05	2.43	0.12
C^e	-	-	-	-	-	-	16.64	0.00

The results for the β' and α vectors are presented in the next sections. The long-run relationships are checked for robustness in Section 5.

3.2.1. The long-run relationships – the bs

Models 1-n and 2-n describe nominal M3 balances of households, while models 3-r and 4-r determine real balances. The models differ in terms of explanatory variables. Table 5 shows the point estimates of the parameters.

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Table 5
The restricted cointegration vectors *ß*

Model 1-n								Test
	m_{t-1}	pc_{t-1}	rc_{t-1}	$rthw_{t-1}$	$(blr - own)_{t-1}$		$GL1_{t-1}$	F(2,59)
	1.000	-1.000	-0.67 [0.04]	-0.67 [0.04]	0.70 [0.09]		-1.23 [0.28]	0.65 [0.53]
Model 2-n	m_{t-1}	pc_{t-1}	rc_{t-1}	rtw_{t-1}	$(blr - own)_{t-1}$		$(p - e)_{t-1}$	F(3,54)
	1.000	-1.000	-0.5 [0.04]	-0.5 [0.04]	0.08 [0.02]		-0.23 [0.02]	1.23 [0.31]
Model 3-r	$(m - pc)_{t-1}$		rdi_{t-1}		$(IRL - IRS)_{t-1}$	dpc_{t-1}	$(p - e)_{t-1}$	-
	1.000		-1.82 [0.016]		0.07 [0.02]	0.23 [0.08]	0.51 [0.08]	-
Model 4-r	$(m - pc)_{t-1}$		rc_{t-1}	C^{e}_{t-1}	$(IRL - OWN)_{t-1}$	dpc_{t-1}	$(p - e)_{t-1}$	F(2,60)
	1.000		-1.000	-0.007 [0.001]	0.08 [0.02]	-	0.46 [0.08]	2.61 [0.08]

Note: α restricted as in Table 6, standard errors in square brackets.

Theoretically, money holdings should be linear homogenous in the price level in the long-run, thus suggesting to impose a parameter restriction of -1 on the long-run parameter for the price level. At the same time, the consumption expenditure deflator used in the empirical analysis might be a restrictive proxy for the price level actually entering households' money holding decisions. In this case parameter estimates larger than one might also be justified. For both nominal models, the assumption of linear homogeneity is not rejected by the data.[15] Linear homogeneity between household balances and prices permit the reformulation of the models in terms of a demand for real money balances. The results of such a reformulation are presented in Annex II (model 1-r), with most features of model 1-n remaining unchanged.[16] This finding provides an empirical justification to impose the restriction from the outset and estimate models 3-r and 4-r in terms of real balances.

Turning to the ***parameter estimates on the scale variables***, i.e. consumption, income and wealth, the following restrictions are proposed:

1. In the case of model 1-n, an over-identifying restriction is introduced by postulating that real consumption and trend housing wealth are equally important for the demand for money, an assumption similar to Thomas (1997). This equality restriction is not rejected. Together, the parameters sum to 1.32, a value not out of line with results from analyses with euro area aggregate M3 (see for instance, Calza et al., 2001). A restriction to the value of unity is however rejected.
2. However, a slightly different constraint implying that the ratio of broad money to consumption is determined by the ratio of total wealth to consumption can be imposed in model 2-n. Indeed, the parameters of real consumption and real total wealth can be constrained to an equal weighting of 0.5 on each variable and thus sum to unity.

[15] The test statistic is distributed as F(1,59) = 0.31 (p-value = 0.58).

[16] Indeed, except for the change in prices which enters the long-run relationship with a negative sign, in-line with an interpretation as an opportunity cost, most features are very similar. This suggests that the deviation from linear homogeneity in model 1 can be considered of second order.

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3. In model 3-r, no over-identifying restrictions are imposed on the cointegration relationship. The parameter estimate obtained for real disposable income of +1.8 is very high, but the model includes only disposable income as a scale variable. A parameter estimate in this order of magnitude is generally thought to reflect the fact that households hold broad money above and beyond transaction purposes, for instance, as a store of wealth.
4. Model 4-r takes a different approach to explain the scale of real household M3 balances. Instead of disposable income, it focuses, like the nominal models, on consumption expenditure. Additionally, expectations of future economic activity (C^e) enter the system. The other variables are similar to model 3-r. A major difference between the two real models is the parameter estimate for real consumption and real disposable income. While the parameter for consumption in model 4-r can be constrained to unity, this is strongly rejected for the parameter of disposable income in model 3-r. Therefore, this parameter restriction seems due to the inclusion of expectations on future economic activity, which substitute for some of the explanatory power of real disposable income in model 3-r.

The parameter estimates of the ***opportunity cost*** variables exhibit the following characteristics:

1. In models 1-n and 2-n, the opportunity costs of holding money are proxied by the spread between a bank lending rate and the own rate of household M3 balances (in logs). In both models, it has the expected sign and is significantly different from zero (at the 5% percent level). However, the parameter estimate in model 2-n is much smaller than in model 1-n (by a factor of 10).
2. In model 3-r, two proxies for opportunity costs of holding money are included: first, the slope of the yield curve is found to have a negative impact on the level of real M3 holdings. Second, consumer price inflation d*pc* has a negative impact on real household money holdings. This is also present in Thomas (1997) and in Coenen and Vega (1999). Restricting the inflation rate parameter to zero is rejected by the data.[17] This finding suggests that inflation is a relevant opportunity cost for households in the long-run, as households shift out of money and into real assets with a higher level of inflation. Comparing the parameters of the slope of the yield and inflation suggests that the substitution between money and real assets may be significantly stronger than between financial assets.
3. By contrast, in model 4-r the parameter on inflation can be constrained to zero, while the parameter estimate for the spread between the long-term bond yield and the own rate on household M3 holdings has a similar magnitude as in model 3-r.

Turning to the variables intended to capture ***precautionary considerations***, the following observations can be made:

1. The sign on the financial market uncertainty measure in the long-run relationship has the expected positive sign and is significantly different from zero, implying that higher financial market uncertainty leads to higher money holdings.
2. In addition, model 1-n also has as an exogenous regressor in the system's short-run dynamics, households' unemployment expectations in the coming twelve months as a proxy for consumer confidence. The point estimate is negative which is in line with Atta-Mensah (2004b) for Canada. Obviously, this reflects the fact that over the sample the effect via precautionary money holdings dominates.
3. In models 2-n, 3-r and 4-r, the price earnings ratio exerts a negative effect on household money holdings, in line with an interpretation that emphasizes the implicit discount factor embodied in it. When corporate earnings relative to observed stock prices are high, money holdings are high owing to the high uncertainty as reflected in the implicit strong discounting of earnings (and vice versa).

[17] $F(1,61) = 11.21$, p-value = 0.00.

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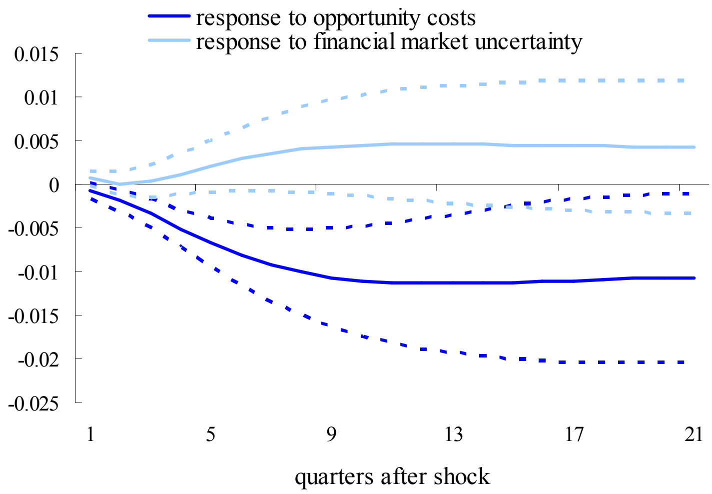
In order to highlight the countervailing impact from portfolio considerations on household money holdings, Figure 1 shows the generalized impulse responses of household M3 to one standard deviation shock in the opportunity cost and uncertainty variables in the context of Model 1-n as an example. A widening of the spread has a significant and negative effect on the level of real M3 holdings by households. In the face of higher borrowing costs, households have an incentive to reduce their holdings of the lower-yielding monetary assets. The complete impact has unfolded after around 10 quarters and remains negative thereafter on the level of money. An increase in the level of financial market uncertainty implies higher money holdings, with the effect taking around 10 quarters to unfold as well. In terms of magnitude, the impact of the interest rate seems to dominate uncertainty effects in this model.

Figure 1
Generalized impulse response of household money to opportunity costs and uncertainty



Note: Dotted lines denote 95% confidence interval around the respective impulse response.

3.2.2. The cointegration relationships – the $\beta' y_t$

The cointegration relationships of all four models are shown in Chart F in annex 1. The charts illustrate quite persistent deviations from the embodied "equilibrium" level. In the case of model 1-n, downside deviations from the average level are observed for periods when the pace of economic activity was slowing (1992–1994, 2001–2003 and since 2007 Q1), while upside deviations are observed particularly for the period 1995–1996 and 1999–2000 before the burst of the dotcom bubble and to a lesser extent between 2004 and 2006.

The co-integration relation from model 2-n exhibits a visibly different pattern from that obtained from model 1-n, especially for the most recent period between 2004 and 2008. The cointegration relationship suggests that M3 holdings have been broadly in line with the level implied by the longer-term determinants for this period and does not point to the sharp decline visible in model 1-n since the end of 2006. A casual inspection of the co-integration relation of model 2-n suggests a break in the series around year 2000. However, checking for parameter stability of the long-run relationship with the Nyblom test and the one-step ahead forecast Chow test does not suggest instability in the parameters of the M3 equation, even in 2007/08 (see Figure 2a). Occasional predictive failures may not be a reason for concern, as these may arise when major shocks occur to the system, while the prediction tests might be useful as a diagnostic tool for parameter stability over a longer time period (Juselius, 2006, p. 164).

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Figure 2a
Chow test for household M3 equation

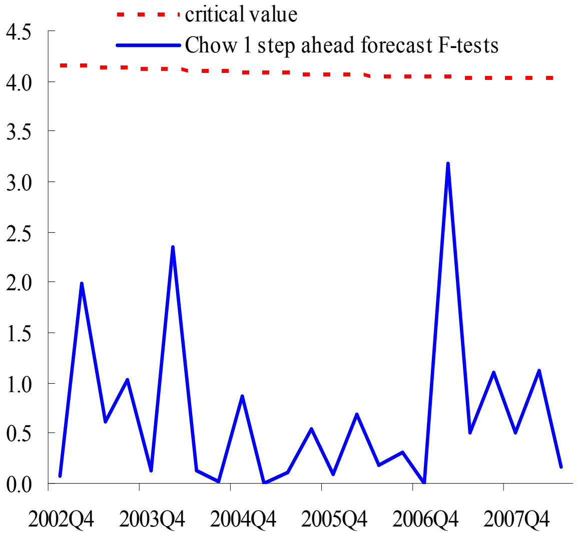
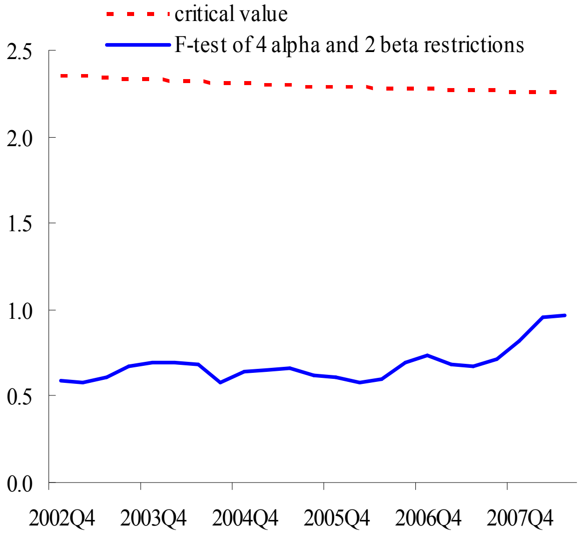


Figure 2b
Recursive F-test on alpha and beta restrictions



Turning to the cointegration relation obtained from model 3-r, it displays certain similarities with that obtained both for model 1-n, with regard to the large positive "peak" in 2000, as well as with that obtained for model 2-n with regard to the assessment of developments in the period 2004–2008. Contrary to both nominal models, model 3-r displays more frequent crossings of the average level. The cointegration relation derived from model 4-r contrasts significantly with that of the other three models. It displays several longer episodes of upward and downward movements. For the more recent period between 2004 and 2008, the assessment of money holdings relative to the long-run determinants would tend to confirm the results obtained from models 2-n and 3-r.

3.2.3. The adjustment to the long-run relationship – the as

With regard to the variables involved in the adjustment to the long-run equilibrium, the tests for weak exogeneity of the variables (presented in Table 4) provide guidance for imposing the exclusion restriction on the a-vector in equation 3.

Table 6
The Loading Factors α

Model 1-n							Test
	Δm_t	Δpc_t	Δrc_t	$\Delta rthw_t$	$\Delta(blr_t - own_t)$	$\Delta GL1_t$	F(4,58)
	-0.0501 [0.009]	-	0.039 [0.009]	-	-	-	1.52 [0.21]
Model 2-n	Δm_t	Δpc_t	Δrc_t	Δrtw_t	$\Delta(blr_t - own_t)$	$\Delta(p\text{-}e)_t$	F(1,54)
	-0.111 [0.009]	-0.046 [0.014]	0.112 [0.022]	0.037 [0.016]	0.628 [0.203]	-	0.08 [0.79]
Model 3-r	$\Delta(m - pc)_t$	Δrdi_t	$\Delta\Delta pc_t$	$\Delta(IRL - IRS)_t$	$\Delta(p - e)_t$		F(3,62)
	-0.026 [0.006]	0.027 [0.006]	-	-	-		1.76 [0.16]
Model 4-r	$\Delta(m - pc)_t$	Δrc_t	$\Delta\Delta pc_t$	$\Delta(IRL - OWN)_t$	$\Delta(p - e)_t$	ΔC^e_t	F(4,61)
	-0.044 [0.009]	-	-	-	-	34.14 [7.85]	0.96 [0.44]

β restricted as in Table 5, standard errors in square brackets.

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The following restrictions on the load factors are compatible with the data:

1. In model 1-n, the test indicates that the load factors on the change in the price deflator, the change in wealth, the interest rate spread and the uncertainty measure can be restricted to zero. This leaves two variables to adjust to disequilibria, *m* and *rc*. The parameters for these two load factors are highly significant, with nominal money and real consumption helping to reduce the disequilibrium in the long-run relationship. A joint test for the restrictions placed on the α-vector cannot be rejected at conventional significance levels (see Table 6). This notwithstanding, the speed of adjustment observed for both variables is rather low, as commonly found in studies of household sector money demand.[18] This renders the short-run dynamics more important. Recursive estimation of the load factors indicates that the parameter estimate has remained unchanged between 2002 Q4 and 2008 Q3, while the same exercise for both *α*- and *ß*-restrictions shows a slight increase since mid-2007, while remaining well below the 5% significance threshold (see Figure 2b).
2. The price earnings ratio is the only weakly exogenous variable in the cointegrated VAR of model 2-n (see Table 4). The parameters for the four remaining load factors are highly significant (at the 5% significance level), with nominal money, real consumption and real wealth helping to reduce the disequilibrium, while again the opposite effect is exerted on the price level (see Table 6).[19] In real terms, however, money still equilibrium corrects.
3. In model 3-r, the tests on the weak exogeneity of the variables suggest that only the yield curve is weakly exogenous (see Table 4). Additional restrictions on the load factors for inflation and the price earnings ratio are not rejected. Therefore, only real money and real disposable income adjust to disequilibria (see Table 6). The speed of adjustment for both variables is highly significant, but very low.
4. The weak exogeneity tests in Table 4 indicate that only real household balances and expectations with regard to economic activity adjust to long-run disequilibria. The t-statistic on the load factor in the household M3 equation is 4.7 and well above the rule of thumb value provided by Juselius (2006). This supports the view that money holdings adjust to imbalances. The tests also suggest that the other variables (except the expectations) are pushing factors for monetary developments (see Table 6). This contrasts with the finding from model 1-n which indicated that real consumption adjusts to monetary disequilibria. However, in model 4-r consumer expectations adjust. Granger causality tests also provide weak evidence that money affects consumption expenditure in this model (p-value = 0.11), while an indirect effect is detected from money to consumer expectations, onto the price earnings ratio (p-value = 0.05) and finally on consumption (p-value = 0.06).

3.2.4. Models' explanatory power for household M3 and misspecification tests

The cointegrated VAR models 1-n and 2-n explain the quarterly changes in households' money balances well, with an adjusted R^2 of 0.69 and 0.73, respectively. The goodness of fit of the equation is also illustrated by Chart D in annex 1 which compares actual and fitted data. The residuals in both models for the household M3 equation show a large spike at the end of 2002 (see Chart E). The cointegrated VAR models incorporating real household balances, model 3-r and 4-r, explain the quarterly changes in households' M3 balances less well than the nominal models. The respective adjusted R^2 is 0.49 and 0.59. Model 2-n also fits the development in real consumption and total wealth surprisingly well with a respective adjusted R^2 of 0.51 and 0.92, while model 3-r is able to explain a noticeable share of the quarterly variation in real disposable income, as evidenced by an adjusted R^2 of 0.42. Model 4-r tracks the quarterly variation in consumer expectations quite well, with an adjusted R^2 of 0.41.

[18] See for instance von Landesberger (2007).

[19] A joint test for all restrictions imposed in model 2-n is not rejected at conventional significance levels [F(4,53) = 0.93, p-value = 0.46]

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Table 7
Residual properties for coinetgrated VAR

Specification test	Test statistic	p-value	Stability test	Test statistic	p-value
Model 1-n					
LM-AR(1)	F(36,51) = 0.85	0.69	LM-PC(3) vs. deterministic variables	1.56	0.67
LM-AR(4)	F(36,48) = 0.88	0.65	LM-PC(3) cointegration	0.90	0.83
Multivariate ARCH	F(441,60) = 1.13	0.28	sup Q(t|T)	1.42	0.70
Normality	F(12,56)= 1.17	0.32	mean Q(t|T)	0.73	0.52
Model 2-n					
LM-AR(1)	F(36,47) = 0.70	0.87	LM-PC(3) vs. deterministic variables	0.76	0.86
LM-AR(4)	F(36,44) = 0.63	0.92	LM-PC(3) vs. cointegration	0.94	0.82
Multivariate ARCH	F(441,53) = 1.05	0.44	sup Q(t|T)	4.54	0.15
Normality	F(12,52) = 2.19	0.03	mean Q(t|T)	0.98	0.29
Model 3-r					
LM-AR(1)	F(25,56) = 1.02	0.46	LM-PC(3) vs. deterministic variables	1.36	0.71
LM-AR(4)	F(25,53) = 1.09	0.39	LM-PC(3) vs. cointegration	1.39	0.71
Multivariate ARCH	F(225,62) = 1.20	0.20	sup Q(t|T)	1.75	0.65
Normality	F(10,60) = 0.38	0.95	mean Q(t|T)	0.85	0.48
Model 4-r					
LM-AR(1)	F(36,54) = 1.17	0.30	LM-PC(3) vs. deterministic variables	0.50	0.92
LM-AR(4)	F(36,51) = 0.69	0.88	LM-PC(3) vs. cointegration	0.52	0.91
Multivariate ARCH	F(441,61) = 1.08	0.37	Nyblom sup Q(t|T)	0.79	0.35
Normality	F(12,59) = 2.79	0.00	Nyblom mean Q(t|T)	0.33	0.30

Notes: LM-AR(1) and LM-AR(4) test statistic calculated as in Johansen (1996). ARCH test follows Warne (2009). Normality test as proposed by Doornik and Hansen (1994). LM-PC(3) tests are based on Teräsvirta (1998) calculated using a third order Taylor expansion. Nyblom sup Q(t|T) and mean Q(t|T) computed as in Hansen and Johansen (1999).

In order to assess the statistical properties of the models, Table 7 reports results from several standard misspecification tests on the residuals of the cointegrated VARs. The misspecification tests indicate absence of autocorrelation of residuals for all the models. Multivariate ARCH effects can also not be detected in the residuals. In the case of models 2-n and 4-r, however, the residuals are not normally distributed. The Nyblom tests conditional on the full sample estimates for the constant and the lagged endogenous parameters do not reveal any instability of the long-run parameters for any of the models. Finally, the LM-tests against the alternative of non-linearity in the deterministic variables or the cointegration parameters, which would capture gradual shifts, also do not suggest parameter non-constancy.[20]

[20] See Teräsvirta (1998). Teräsvirta and Eliasson (2001) investigate non-linearity in an error correction model of UK money demand.

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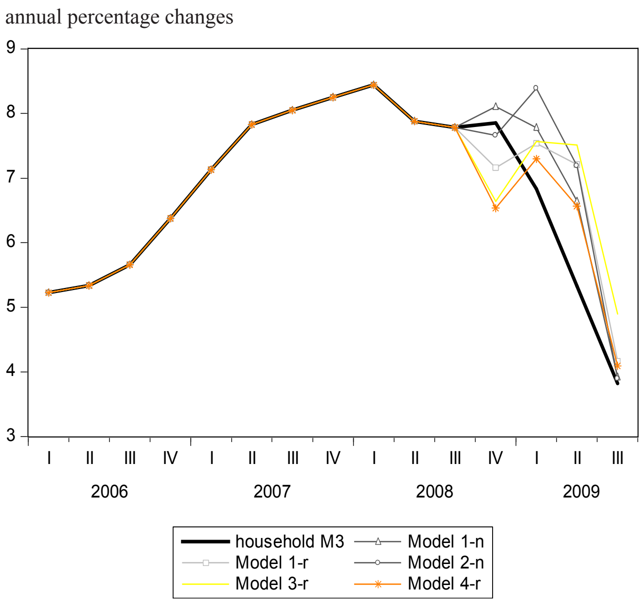
4. EVALUATING THE FORECASTING PERFORMANCE OF THE MONEY DEMAND SYSTEMS

The results presented in Section 3 suggest that the four models describe money demand by euro area households in a satisfactory manner, when judged, for instance, by the in-sample fit and standard misspecification tests. In addition, the estimates for the parameters allow for a theory-consistent interpretation and thereby support the view that money demand relationships have been identified. However, in order to gain additional insights on the models' ability to explain monetary developments, the last four available observations (2008 Q4 to 2009 Q3) are used to produce out-of-sample forecasts. The period covers the intensification of the financial turmoil following the default of Lehman Brothers, which has proven to be challenging for empirical models. In this context, it should be borne in mind that complicated models may have more explanatory power in sample, but also tend to include more variables that can lead to bad forecast results when changes to the economic environment occur. Thus, a more parsimonious specification may be advantageous.

Figure 3 illustrates the forecasts in terms of annual growth rates of household M3 using the actual observations for the other explanatory variables. It suggests that the nominal models and in particular model 1-n predicts monetary developments quite well, capturing the overall pattern of the slowdown, while the predictions from the real specifications suffer from the lack of capturing the rapid slowdown in price developments. More specifically, the strength of the slowdown in household M3 growth in 2009 Q2 was not captured by the models in a convincing manner, while most models do produce a prediction close to the actual outcome for 2009 Q3.

Figure 3
Forecast performance



Notes: for the period 2008 Q3–2009 Q3 actual observations for other explanatory variables.

In short forecasting samples characterized by structural breaks, cointegration models may not be able to exploit the advantage of having an identified long-run relationship. A different perspective on the ability of the cointegrated VAR systems to explain monetary developments is obtained when simulating out-of-sample the money growth for an extended period. Figure 4 shows the outcome of such an exercise conducted with model 2-n and 4-r from 1999 Q1 onward. The forecast for real household M3 growth settles at a stable annual steady state growth rate of 2.5% within eight quarters for model 4-r and takes twice as long for model 2-n. Investigating alternative

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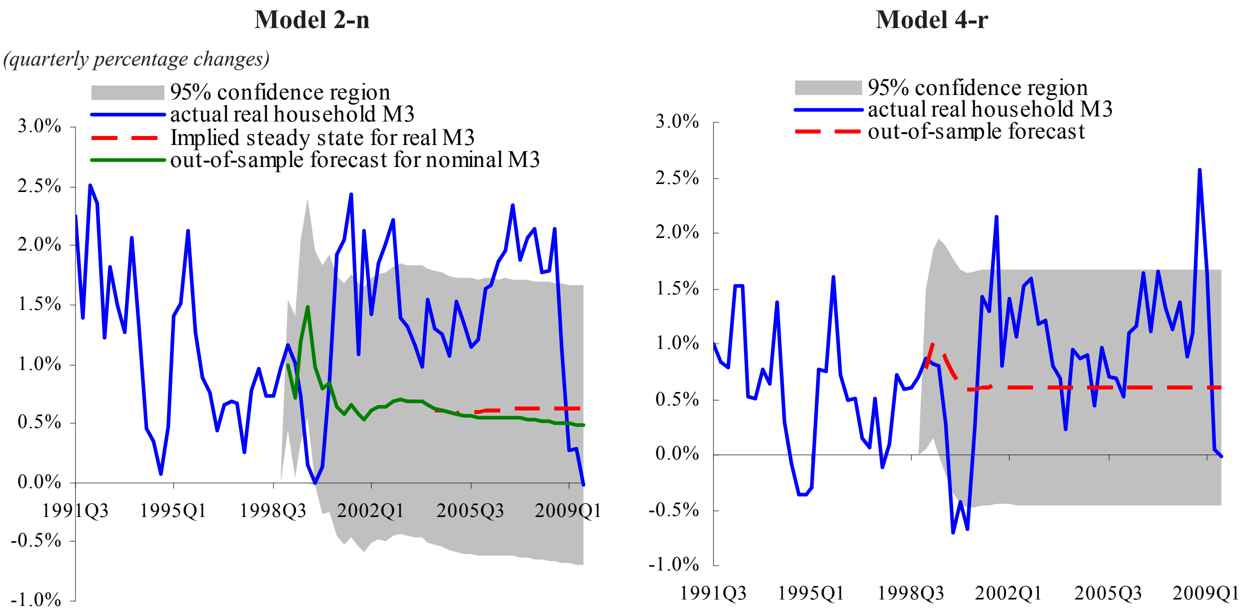
forecast horizons provides similar steady state growth rates. The simulations support the ECB's assessment that a number of exceptional shocks have hit euro area monetary developments over the past ten years, evidenced by the fact that actual M3 growth leaves the 95% confidence region: in 2000 in the context of the dot-com bubble with very low money growth, and later with very high money growth in 2001 Q3 owing to 11 September 2001 and in 2008 Q4 after the default of Lehman Brothers.[21] However, the models also clearly illustrate that household M3 growth has exhibited protracted periods of above steady-state growth, between 2002 and 2004 linked to exceptional portfolio shifts into money, and between 2006 and 2008, as money growth has been boosted by rapid growth of loans for house purchases in the euro area.

Figure 4
Out-of-sample forecast performance 1999:1–2009:3



5. CROSS-CHECKING WITH OTHER ESTIMATION METHODS

A repeatedly voiced observation with regard to the standard cointegrated VAR methodology is that the parameters of interest in the long-run relationship may be affected by the inclusion in the VAR set-up of less relevant variables. In order to cross-check the results obtained with the Johansen methodology, an alternative estimation is conducted using Fully Modified-OLS proposed by Phillips and Hansen (1990). This is a single equation regression with non-stationary variables. In the presence of several model variables affected by the long-run relationships, i.e. not all variables are weakly exogenous, the FM-OLS estimator will not be efficient as the move to the single equation neglects relevant information that could lead to a better point estimate. Nonetheless, if the residuals of the jointly error correcting variables are uncorrelated, this may be a restrictive yet informative exercise. For this purpose, the constrained specifications are re-estimated. The outcome of the exercise is provided in Table 8 showing the point estimates for the parameters of the cointegration relationship from the cointegrated VAR and the FM-OLS procedure.

[21] European Central Bank (2005, 2007).

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Table 8
Estimation results with FM-OLS

	Model 1-n			Model 2-n			Model 3-r			Model 4-r		
Parameter estimate on	Coint VAR	FM-OLS	*Bootstrapped interval 10%*	Coint VAR	FM-OLS	*Bootstrapped interval 10%*	Coint VAR	FM-OLS	*Bootstrapped interval 10%*	Coint VAR	FM-OLS	*Bootstrapped interval 10%*
m	1 -	1	-	1	1	-	-	-	-	-	-	-
m-pc	-		-	-	-	-	1	1	-	1	1	-
pc	+1	+1	-	+1	+1	-	-	-	-	-	-	-
rc	+0.66 [0.04]	+0.50 [0.08]	*+0.95 +0.48*	+0.5	+0.42 [0.10]	*0.58 0.42*	-	-	-	+1	+1.14 [0.01]	-
rdi	-	-	-	-	-	-	+1.82 [0.16]	+1.12 [0.01]	*+3.47 -0.67*	-	-	-
rtw	-	-	-	+0.50	+0.55 [0.08]	*0.58 0.42*	-	-	-	-	-	-
rthw	+0.66 [0.04]	+0.50 [0.07]	*+0.95 +0.48*	-	-	-	-	-	-	-	-	-
dpc	-	-	-	-	-	-	-0.23 [0.08]	-0.06 [0.04]	*+0.67 -1.59*	0 -	-	-
blr-own	-0.67 [0.10]	-0.07 [0.04]	*-0.34 -1.56*	-0.07 [0.03]	-0.06 [0.04]	*-0.01 -0.16*	-	-	-	-	-	-
IRL-IRS	-	-	-	-	-	-	-0.07 [0.02]	-0.00 [0.01]	*+0.19 -0.42*	-	-	-
IRL-OWN	-	-	-	-	-	-	-	-	-	-0.08 [0.02]	-0.04 [0.01]	*-0.05 -0.16*
GL1	+1.20 [0.29]	-0.03 [0.17]	*+0.42 +1.99*	-	-	-	-	-	-	-	-	-
p-e	-	-	-	-0.23 [0.02]	-0.07 [0.03]	*-0.18 -0.29*	-0.51 [0.08]	-0.17 [0.04]	*+0.94 -2.33*	-0.45 [0.08]	-0.18 [0.04]	*-0.27 -0.92*
C_EXP	-	-	-	-	-	-	-	-	-	+0.007 [0.001]	0.00 [0.00]	*+0.00 +0.02*
Equality of models F-Test			F(5,60) = 6.8580 [0.03]			F(5,54) = 8.1874 [0.00]			F(4,62) = 8.7712 [0.00]			F(5,61) = 6.7599 [0.02]

Note: Standard errors in square brackets below coefficients.

The results suggest that the individual point estimates obtained by both econometric techniques are fairly similar. In addition, the table also reports the interval obtained from bootstrapping the β-estimates and imposing a 10% confidence interval.[22] In this respect a number of points are worth mentioning:

[22] A higher confidence level such as 5% would have increased the width of the confidence bands significantly.

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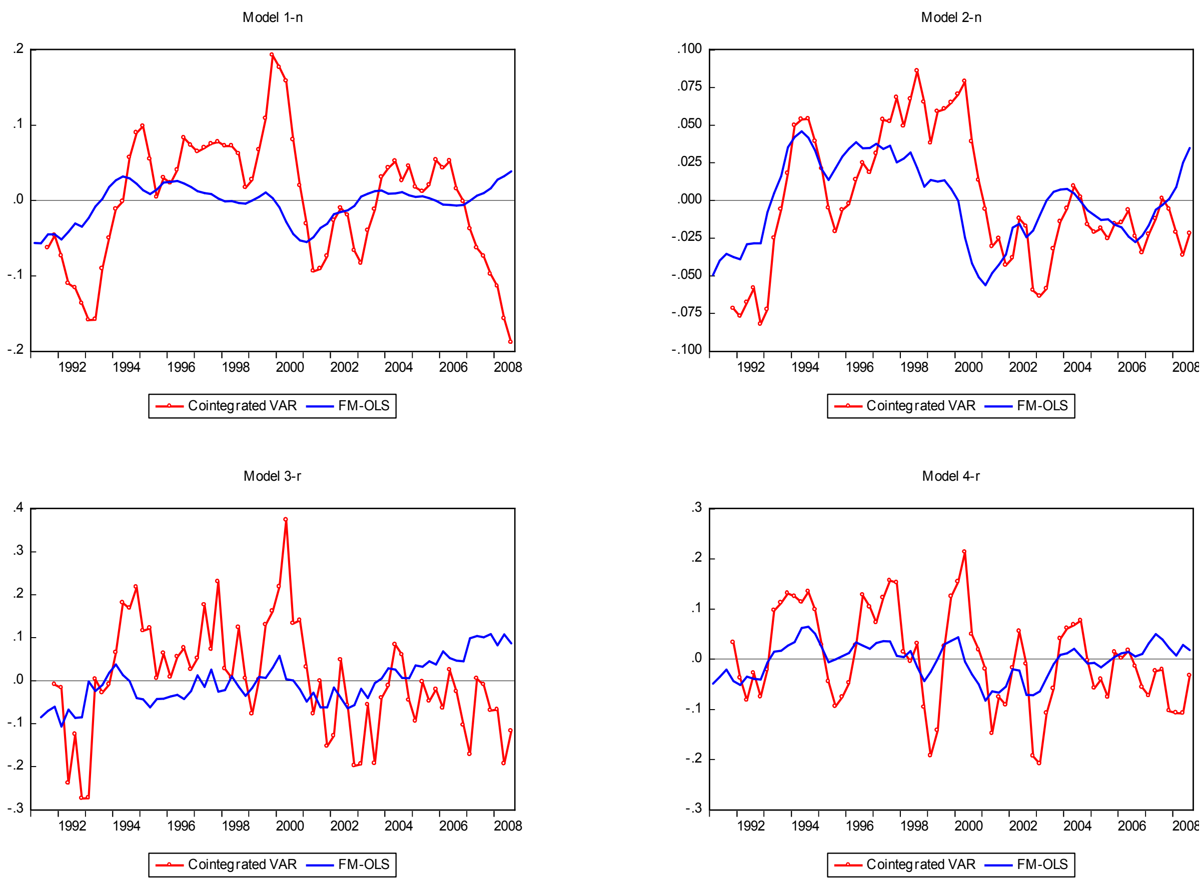
- The parameters of models 2 and 3 are included in the bootstrap interval, which are, however, excessively wide.
- The parameter estimates obtained with FM-OLS for the scale variables (consumption, disposable income, wealth) are generally included in the bootstrapped intervals of the cointegrated VAR.
- The outcome is more mixed for the other explanatory variables (opportunity costs, uncertainty related variables). The parameter estimates on these other variables, if significant in the FM-OLS equation, tend to be different from those obtained in the cointegrated VAR implying a distinctive assessment of the importance of these variables for money demand.
- Jointly imposing the parameter estimates obtained from the FM-OLS procedure in the original cointegrated VAR framework leads to a rejection of the equality of the estimates at the 5% level in all cases (see the last row with p-values in brackets).

However, gauging the similarity of the cointegration relationships by the cross correlation between their monetary overhang measures suggests that models 2-n and 4-r are quite similar in their assessment of actual money holdings relative to equilibrium household M3 holdings with contemporaneous correlation coefficients around 0.55 for the period 1991 Q1 to 2008 Q3. By contrast, for models 1-n and 3-r a similarity between the two measures cannot be found contemporaneously (0.19 and 0.00). Cross-correlation analysis suggests that the Johansen measures tend to lead their respective FMOLS measures with a slightly better match. Figure 5 shows the monetary overhangs from the cointegrated VAR (red line) and FM-OLS estimation (blue line) of all models.

Figure 5
Comparison of monetary overhang measures



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Monetary overhangs are often considered as an appropriate metric for measuring excess liquidity. When gauging the similarity of the assessment by whether both series are over or under the zero line in Figure 3, the superiority of models 2-n and 4-r is again clearly visible.

The chart also illustrates the different development in the estimated monetary imbalances during the financial turmoil from mid-2007 onwards. In this period a sharp decrease in the opportunity costs of holding money and an increase in uncertainty was observed. It is not surprising that the monetary overhang measures estimated using FM-OLS that downplay these factors differ especially in this period.

6. SUMMARY AND CONCLUSION

In the euro area, household holdings of M3 have been found to be informative with regard to developments in HICP inflation. An empirical framework permitting to analyse the driving factors for household money demand is therefore an important element for monetary analysis. The paper presented several different approaches to model the demand for nominal and real household M3 balances in the euro area. In investigating the long-run relationship between money, different scale variables and opportunity costs, only a few combinations may satisfy formal cointegration tests, even if an underlying cointegration relationship is present for a broader set of similar variables (see Ericsson, 1998). Several important outcomes have been found.

Neither nominal models rejects linear homogeneity between money balances and the price level. They therefore support the specification in real terms and suggest that in the long-run households are not subject to money illusion, in line with theoretical considerations.

1. Household money balances are never weakly exogenous with regard to the other variables of the cointegrated VARs and therefore always adjust to disequilibria between (real) money and its long-run determinants. That said, the models also provide evidence that the volume of transactions (proxied by real disposable income or real consumption) is affected by the monetary disequilibria and also adjusts. By contrast, measures of wealth, opportunity costs and financial market uncertainty are generally found to be the forces jointly determining the growth of money and real income/real consumption in the long-run.
2. In explaining households' broad money balances, wealth, and in particular housing wealth, is found to play an important role.[23] However, it seems to be wealth in conjunction with either real consumption expenditures or real disposable income that best captures households' notional level of money holdings. Omitting wealth from the specification leads to a sizeable increase in the income elasticity of money demand. At the same time, the inclusion of consumer expectations with regard to economic activity is able to offset this increase. This may reflect the fact that, in theory, wealth captures expectations on the future income path.
3. Interest rate developments seem to play a significant role for the development of household balances. Models specified with a double-log formulation for opportunity costs exhibit a markedly stronger impact than is the case for the semi-log functional forms. However, in all models, an increase in opportunity costs leads to a significant decline in (real) money holdings with the effect fully materialising after about 10 quarters.
4. The different models suggest that the impact of uncertainty on household balances is complex. Financial market uncertainty clearly stimulates M3. By contrast, economic uncertainty exhibits a more ambiguous impact on money holdings reflecting, on the one hand, the boosting impact stemming from the precautionary motive and, on the other, the dampening impact from a transactions motive.

23 See de Bondt (2009) and Beyer (2009) for a similar finding with regard to aggregate M3.

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5. Correctly incorporating the persistent behaviour of interest rates and uncertainty into the money demand function is essential to adequately capture the driving forces impacting on money and expenditures as well as their mutual interaction. The different specifications presented suggest that there are several modelling approaches that can be undertaken.
6. All models pass a battery of misspecification and stability tests. Moreover, the parameter estimates are checked with a (probably more robust) single-equation approach. Especially at the end of the sample, differences in the models' estimates are obvious.

While the outcome of the exploration may not be seen as surprising as the estimates are consistent with results reported in the literature, the exercises presented help to better identify the determinants of euro area money demand and to interpret current monetary developments. As households' money demand captures the bulk of aggregate euro area M3, it should also be helpful in understanding the long-run money-price-nexus.

More generally, the exercise also provides insights that go beyond the portfolio allocation decision of households. According to our analysis, it is quite apparent that in equilibrium, households jointly determine consumption and broad money holdings, influenced by both wealth as well as interest rates. The importance of household money holdings for consumption expenditures may cast doubt on a purely passive role for money in this context. Moreover, as both bank lending rates and the own rate of households M3 are found significant, the determination of money holdings seems to interact with wealth and borrowing. In order to be able to analyse the interaction between money holdings, consumption and wealth more fully, the financing of households needs to be modelled as well, which goes beyond the scope of this paper.

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ANNEX 1

Table 9
Results of unit root tests

Variables	(D,X)	ADF t-Statistic	ADF p-value*		KPSS LM-Statistic	critical value 10%	critical value 5%
level						**10%**	**5%**
m	(CT,6)	-0.98	0.94	(CT)	0.6938	0.119	0.146
pc	(CT,9)	-1.14	0.92	(CT)	0.1049	0.119	0.146
rtw	(CT,12)	-1.75	0.72	(CT)	0.2358	0.119	0.146
rthw	(CT,12)	-1.88	0.65	(CT)	0.1908	0.119	0.146
rc	(CT,10)	-0.84	0.96	(CT)	0.1349	0.119	0.146
rdi	(CT,11)	-1.93	0.63	(CT)	0.1006	0.119	0.146
C^e	(C,11)	-2.19	0.21	(C)	0.1999	0.347	0.463
UN^e	(C,4)	-3.03	0.04	(C)	0.3467	0.347	0.463
GL1	(C,12)	-0.88	0.79	(C)	0.3903	0.347	0.463
p-e	(C,3)	-1.33	0.61	(C)	0.1497	0.347	0.463
IRL-IRS	(C,11)	-3.95	0.00	(C)	0.2156	0.347	0.463
blr-own	(C,1)	-1.46	0.55	(C)	0.3465	0.347	0.463
IRL-OWN	(C,10)	-2.61	0.10	(C)	0.4503	0.347	0.463
1st difference							
m	(C,0)	-3.45	0.01	(C)	0.2072	0.347	0.463
pc	(C,2)	-3.55	0.01	(C)	0.3393	0.347	0.463
rtw	(C,3)	-2.09	0.25	(C)	0.1928	0.347	0.463
rthw	(C,11)	-2.79	0.07	(C)	0.3706	0.347	0.463
rc	(C,9)	-2.30	0.17	(C)	0.1827	0.347	0.463
rdi	(C,10)	-1.94	0.31	(C)	0.1834	0.347	0.463
C^e	(C,10)	-3.40	0.00	(C)	0.1232	0.347	0.463
UN^e	(C,4)	-3.64	0.00	(C)	0.0510	0.347	0.463
GL1	(N,11)	-3.88	0.00	(C)	0.0882	0.119	0.146
p-e	(N,0)	-6.87	0.00	(C)	0.3404	0.347	0.463
IRL-IRS	(N,0)	-5.54	0.00	(C)	0.1625	0.347	0.463
blr-own	(N,2)	-4.70	0.00	(C)	0.2546	0.347	0.463
IRL-OWN	(N,9)	-4.18	0.00	(C)	0.0465	0.347	0.463

Note: ADF-test: with MacKinnon (1996) one-sided p-values, KPSS: Kwiatkowski-Phillips-Schmidt-Shin (1992, Table 1). (D,X) with D indicating that the estimated regression includes the following deterministic terms: C – constant, CT – constant and trend, N – no deterministic terms. X indicates the number of lagged endogenous terms retained in the estimated test regression (with at least 5% significance) starting from a maximum of 12 lags. Cut-off is determined by sequential testing on the t-statistic of the lagged endogenous variables with at least 5% significance level. KPSS test using Bartlett kernel with cut-off determined by automatic Andrews (1991) procedure.

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Table 10
Lag length determination for VAR

Model	Criterion	Lag length				
		0	1	2	3	4
1-n	Likelihood Ratio Test	NA	2186.05	119.68*	43.86	43.16
	Akaike Information Criterion	-24.03	-44.46	-45.74*	-45.58	45.39
2-n	Likelihood Ratio Test	NA	1330.82	217.75	109.98*	31.22
	Akaike Information Criterion	-19.00	-40.11	-43.07	-44.29*	-43.96
3-r	Likelihood Ratio Test	NA	1484.53	55.73*	28.34	22.46
	Akaike Information Criterion	-3.21	-27.20	-27.46*	-27.27	-27.01
4-r	Likelihood Ratio Test	NA	937.35	74.14*	36.97	29.26
	Akaike Information Criterion	-7.34	-22.13	-22.45*	-22.14	-21.76

Table 11
Tests for stationarity of variables

	Null Hypothesis: variable k is stationary Alternative hypothesis: variable k is not stationary							
	Model 1-n		**Model 2-n**		**Model 3-r**		**Model 4-r**	
Equation for	F(5,59)	p-value	F(5,53)	p-value	F(1,60)	p-value	F(5,60)	p-value
m	7.49	0.03	9.68	0.00	-	-	-	-
m-pc	-	-	-	-	8.86	0.01	7.98	0.01
pc	7.44	0.03	9.40	0.00	-	-	-	-
rc	7.64	0.03	8.99	0.00	-	-	7.55	0.02
rdi	-	-	-	-	8.34	0.01	-	-
rtw	-	-	9.38	0.00	-	-	-	-
rthw	7.43	0.05	-	-	-	-	-	-
dpc	-	-	-	-	4.78	0.03	5.08	0.03
blr-own	6.51	0.04	8.30	0.00	-	-	-	-
IRL-IRS	-	-	-	-	6.97	0.01	-	-
IRL-OWN	-	-	-	-	-	-	7.30	0.01
GL1	7.42	0.02	-	-	-	-	-	-
p-e	-	-	7.55	0.00	5.47	0.02	5.13	0.03
C^e	-	-	-	-	-	-	3.18	0.11

Note: p-values derived from comparison with respective bootstrapped distributions.

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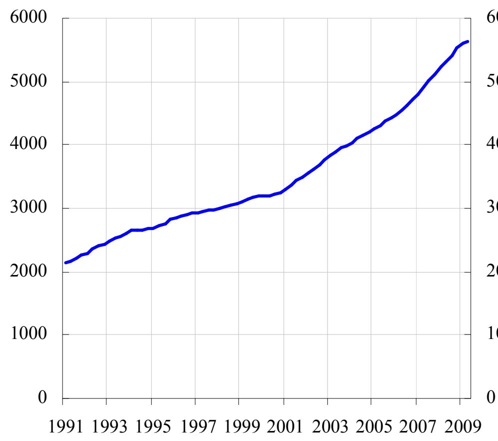
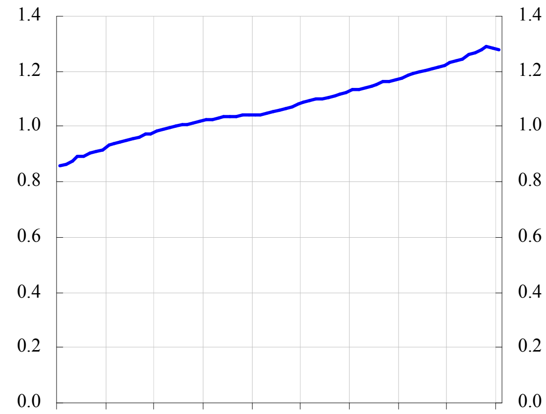
Chart A
Main variables used in the cointegrated VAR systems

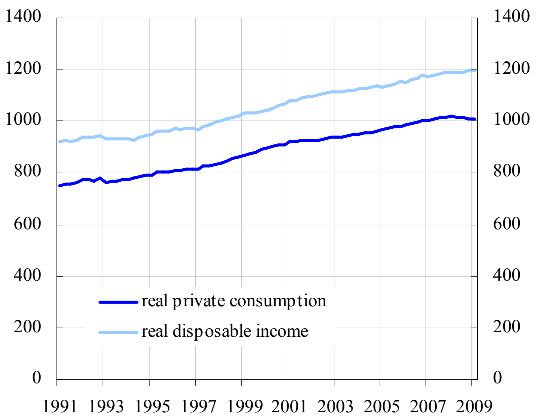
household M3 holdings (in EUR bn)



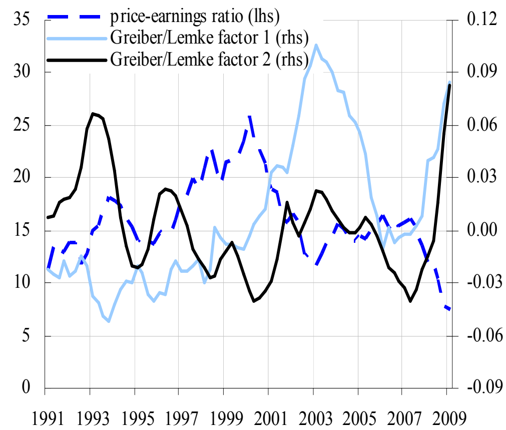
deflator of private consumption



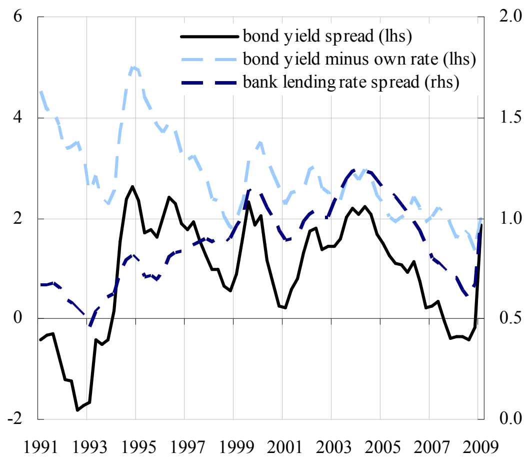
real private consumption and real disposable income (in EUR bn)



price- earnings ratio & Greiber/Lembke uncertainty measures



Interest rate measures (in p.p.)



Real household wealth (in EUR bn)

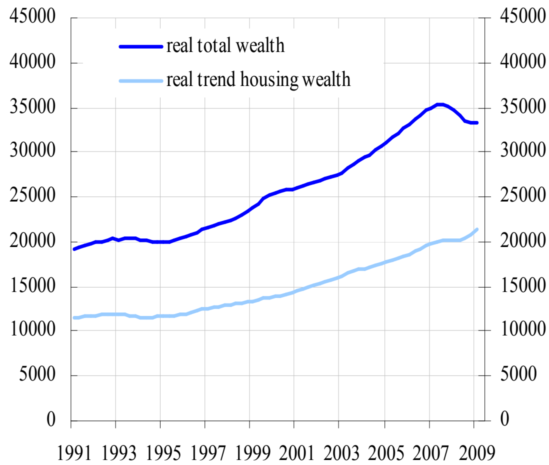


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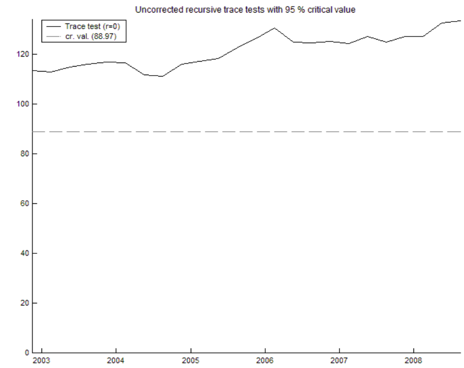
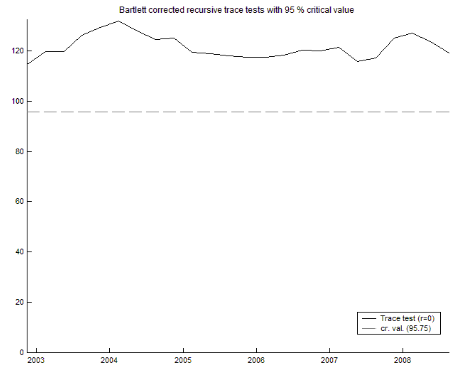
Chart B
Recursive trace test results

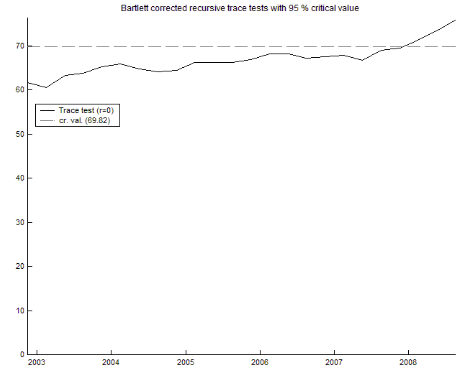
Model 1-n



Model 2-n



Model 3-r



Model 4-r

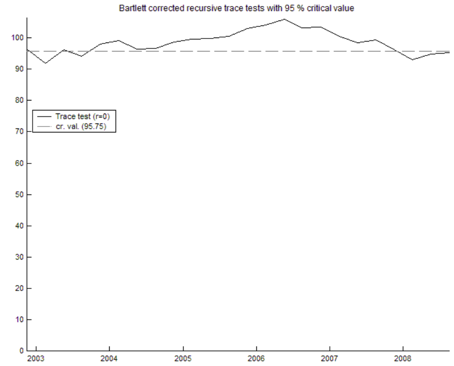
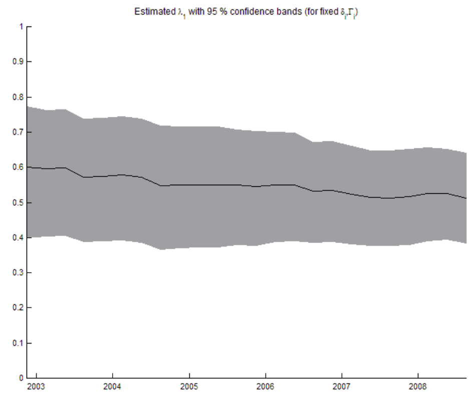
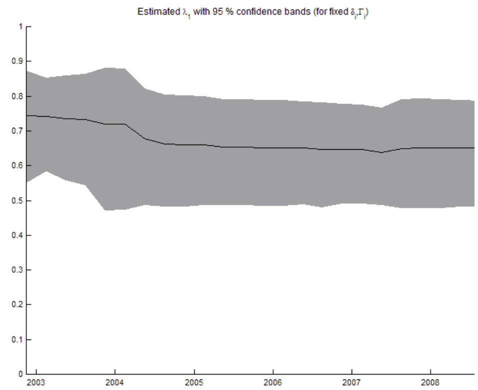


Chart C
Estimated λ_1 with 95% confidence bands (for fixed short run dynamics)

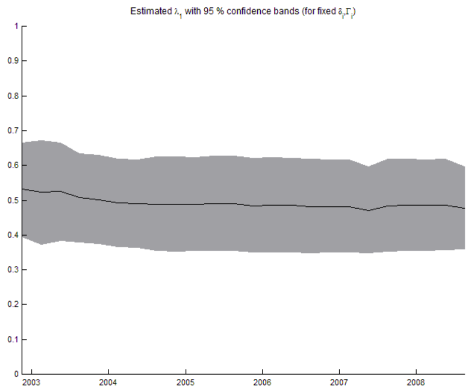
Model 1-n



Model 2-n



Model 3-r



Model 4-r

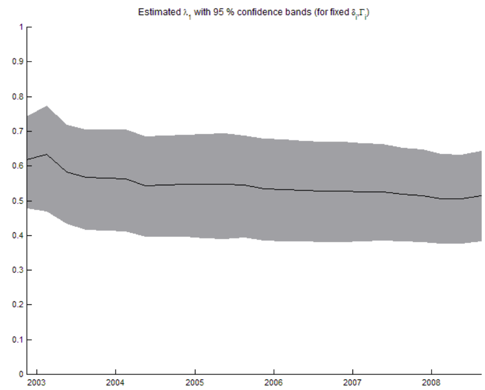


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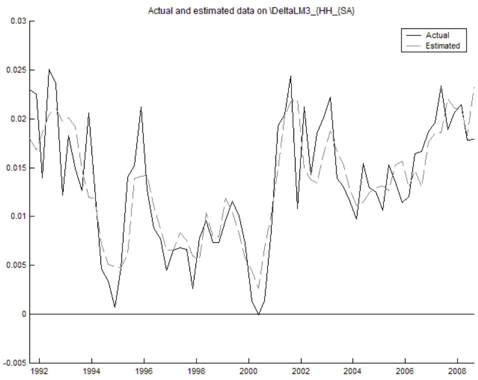
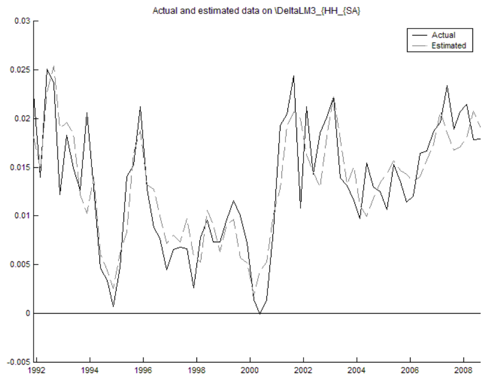
Chart D
Actual & fitted

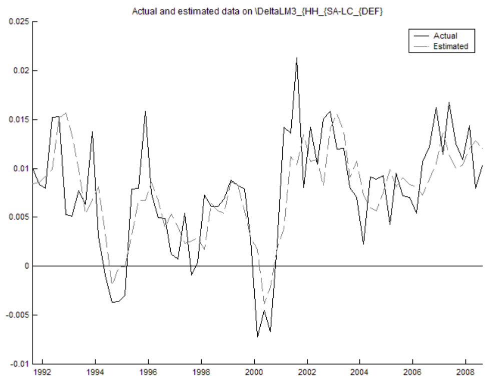
1-n



2-n



3-r



4-r

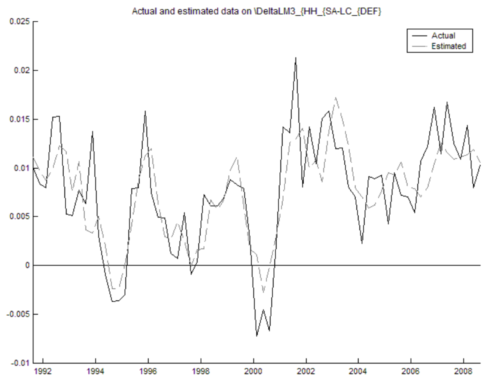
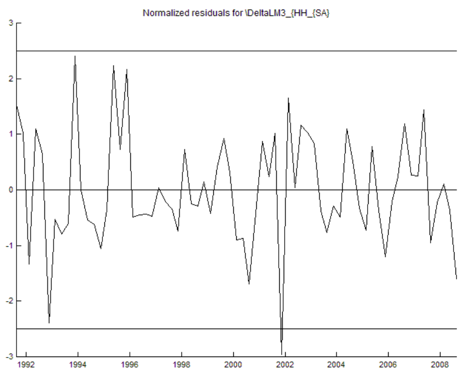
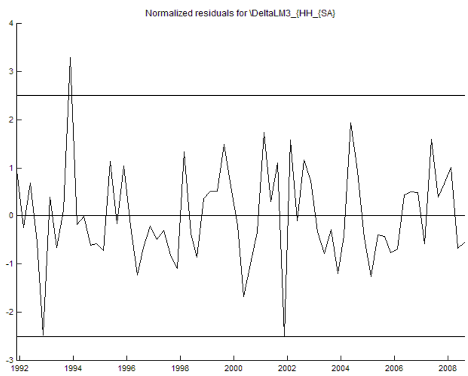


Chart E
Normalised residuals

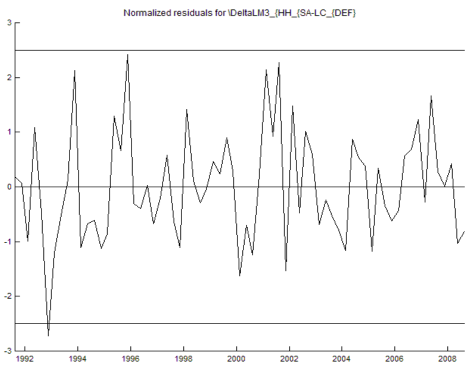
1-n



2-n



3-r



4-r

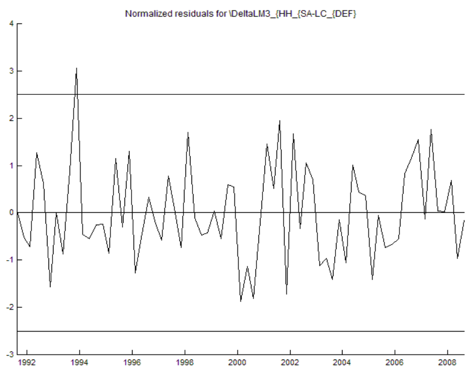


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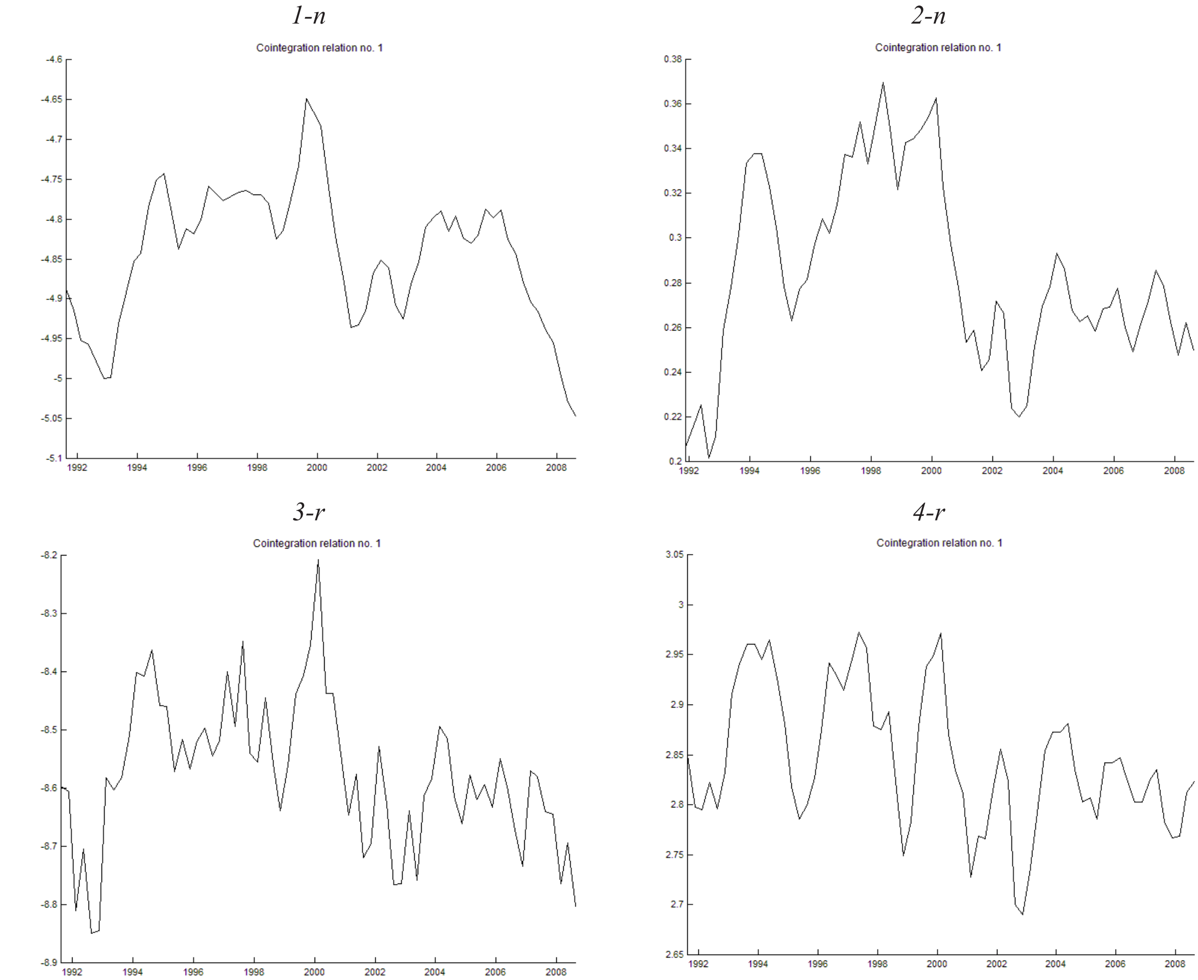
Chart F
Cointegration relationships



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ANNEX 2
Model 1-r

Table 12
Lag length selection for model 1-r

Lag length	**0**	**1**	**2**	**3**	**4**
Likelihood Ratio Test	NA	1089.94	125.62*	30.85	48.57
Akaike Information Criterion	-26.48	-44.18	-45.51*	-45.08	45.21

Table 13
Residual properties for model 1-r

Model 1-r					
LM-AR(1)	F(36,48) = 1.09	0.38	Multivariate ARCH	F(441,61) = 1.19	0.20
LM-AR(4)	F(36,51) = 1.14	0.33	Normality	F(12,59) = 0.85	0.60
For rank 1					
Nyblom Sup F		18.97			0.46
Nyblom Mean		5.49			0.50
PKK S(9)		1.32			0.39

Table 14
Trace test results for model 1-r

1-r	test	124.59	67.98	39.25	13.97	6.20	1.44	-
	p-value	0.000	0.000	0.259	/	/	0.230	-
	bootstrapped p-value	0.048	0.619	0.717	0.950	0.657	0.449	-

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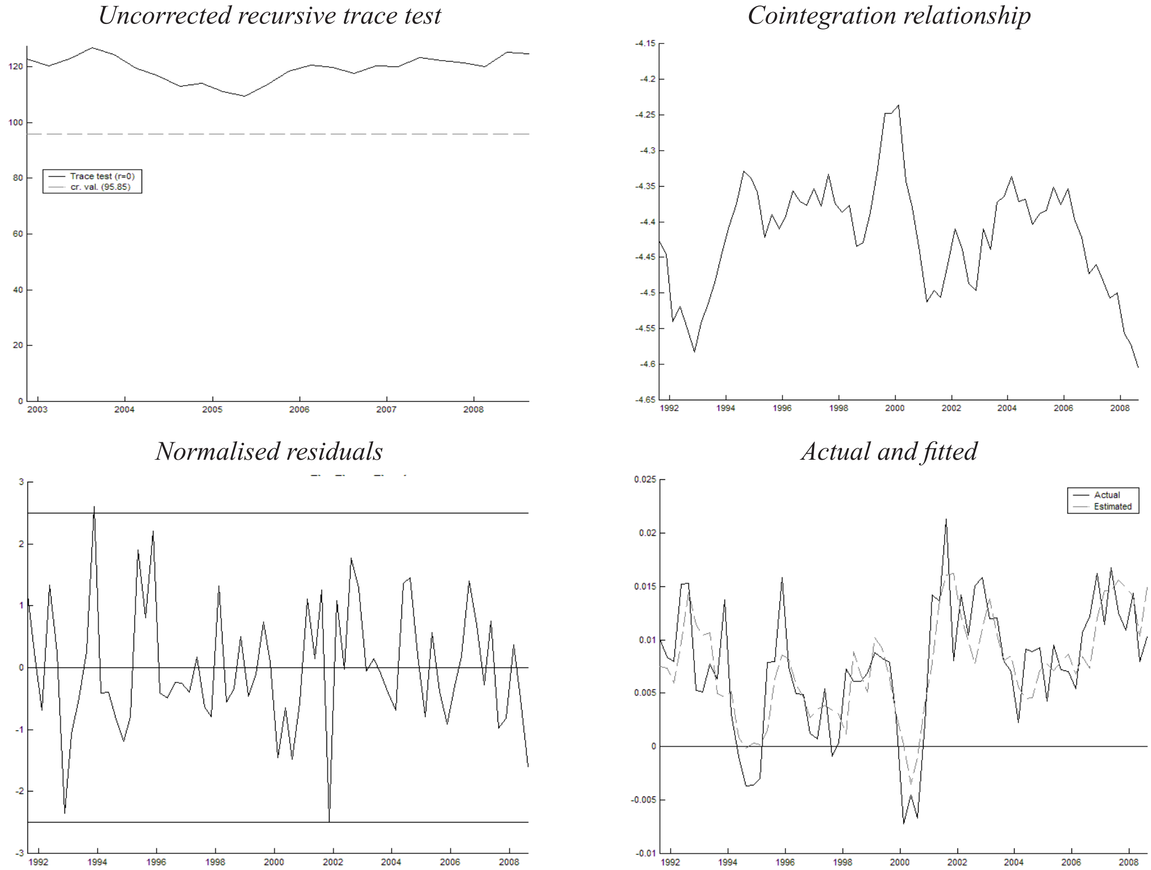
Chart G
Charts for model 1-r



Long-run relationship for model 1-r

$$\begin{bmatrix} \Delta(m-pc)_t \\ \Delta rc_t \\ \Delta rthw_t \\ dpc_t \\ \Delta(blr_t - own_t) \\ \Delta GL1_t \end{bmatrix} = \begin{bmatrix} \underset{0.008}{-0.044} \\ - \\ \underset{0.009}{0.038} \\ \underset{0.005}{0.011} \\ - \\ - \end{bmatrix} \left[(m-pc)_{t-1} - 0.50 rc_{t-1} - \underset{0.04}{0.74} rthw_{t-1} + \underset{0.09}{0.77} (blr-own)_{t-1} - \underset{0.22}{1.29} GL1_{t-1} + \underset{3.01}{7.00} dpc_{t-1} \right] + \ldots$$

Table 15
Residual properties for Model 1-r

Specification test	Test statistic	p-value	Stability tests	Test statistic	p-value
LM-AR(1)	F(36,53) = 0.76	0.81	LM-PC(18) lagged endogenous	10.51	0.91
LM-AR(4)	F(36,50) = 0.68	0.89	LM-PC(3) cointegration	1.22	0.75
Multivariate ARCH	F(441,60) = 1.09	0.35	Nyblom sup Q(t|T)	1.9315	0.91
Normality	F(12,58) = 0.58	0.85	Nyblom mean Q(t|T)	0.9648	0.82

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Feedback to the ECB's Monetary Analysis: The Bank of Russia's Experience with Some Key Tools*

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ABSTARCT

The paper investigates to what extent some basic tools of the ECBs monetary analysis can be useful for other central banks given their specific institutional, economic and financial environment. We take the case of the Bank of Russia in order to show how to adjust methods and techniques of monetary analysis for an economy that differs from the euro area as regards, for instance, the role of the exchange rate, the impact of dollarization and the functioning of sovereign wealth funds. A special focus of the analysis is the estimation of money demand functions for different monetary aggregates. The results suggest that there are stable relationships with respect to income and wealth and to a lesser extent to uncertainty variables and opportunity costs. Furthermore, the analysis also delivers preliminary results of the information content of money for inflation and for real economic development.

JEL classification: E41, E52, E58

Keywords: Money demand, transition countries, cointegration analysis, inflation, real economic activity

1. INTRODUCTION

Monetary analysis at central banks has different meanings across the world and over time. Some parts of the world may still focus on quantitative targets for (base) money and thereby

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blur the meaning of operational and intermediate targets and indicators or reference values. In contrast, in its two-pillar strategy, the ECB makes it clear that it uses the monetary pillar to collect information on medium- to long-term risks to price stability by focusing on the analysis of money and credit aggregates. It thus ensures a "full information approach" that may otherwise be dominated by the analysis of cyclical movements of the economy and the information on short-term risks[2]. Monetary analysis at the ECB has been an evolutionary process during which tools and techniques have developed as described in Papademos and Stark (2010). This process has been monitored by other central banks which set up new strategies for an autonomous monetary policy that focuses on internal price stability rather than on stable exchange rates. We describe the Bank of Russia's experience in this respect and to what extent some key tools of monetary analysis as practiced by the ECB can be useful for it. On the one hand, the Bank of Russia may benefit from tools that are already regularly used in the ECB's monetary assessment. The composition of drivers behind money-stock growth indicates that the Russian economy is evidently prone to exogenous money-supply shocks. Identifying these shocks and their macroeconomic consequences is an important practical task for day-to-day monetary policy analysis. The models developed to interpret monetary developments which constitute an essential part of the ECB's monetary analysis seem particularly suitable for this task. On the other hand, simply copying the tools would not be advisable as the economic and financial environment in Russia differs to some extent from the euro area. In both, their financial sectors have in common the fact that they are rather bank-based than capital-market-based. Financial markets in Russia, however, are less deep and less liquid compared to the euro area and money might be the most important financial store of value for a large proportion of the population. Furthermore, high inflationary and hyperinflationary periods are closer in the collective memory than in the euro area and foreign currency has often served as a safe haven. Currency substitution, or, in its broader definition, "dollarization" has inertia and monetary aggregates that include foreign denominated components should behave differently to those that do not. External nominal anchors have dominated monetary policy in the past and exchange-rate developments have triggered rapid reactions of money holders. Last but not least, Russia is an oil-exporting economy and sovereign wealth funds help to buffer the impact of commodity-price fluctuations and to save financial resources for future generations during normal times. During turbulent times they can also function as crisis tools and provide additional funding. Their behavior can significantly influence money creation and thereby, may be understood as exogenous factors or supply-side factors which influence monetary developments beyond the usual money demand factors.

We acknowledge these differences in our study and focus on some key tools of ECB monetary analysis as described in chapters 3 and 4 of Papademos and Stark (2010) which we apply to the Russian case. We start with a brief review of the role of money in the Bank of Russia's monetary policy since the early 1990s and a description of monetary developments, given in sections 2 and 3. Section 4 forms the core of the paper, as it presents money demand estimations for different monetary aggregates. In section 5 we analyze the information content of money for inflation and real economic activity and in section 6 we conclude.

2. THE ROLE OF MONEY IN THE BANK OF RUSSIA'S MONETARY POLICY – A REVIEW

The main stages of evolution of the conduct of monetary analysis and its role in the Bank of Russia's (CBR) monetary policy framework may be provisionally described by considering five different periods. They highlight the role of money in an economic environment which suffered

[2] Notably, after the Global Financial Crisis the role of monetary analysis is also emphasized in connection with financial stability objective and macroprudential policy. The analysis of this issue however remain beyond the scope of this paper.

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from periods of price and financial instability and shifted from a fixed to a managed exchange rate regime.

The early 1990s. The CBR paid serious attention to monetary analysis and the developments of monetary aggregates as soon as the first steps to liberalize the economy were taken in the early 1990s. The transition from a planned to a market economy caused drastic structural shifts in both the real and the financial sector in Russia. In these circumstances the CBR's monetary policy was conducted against the background of the hyperinflation that followed the lifting of price regulation, deep recession of the real sector, depreciation of the national currency and high macroeconomic uncertainty. The CBR had to find a balance between restraining inflation and supportive measures aimed at preventing the collapse of the real economy and the domestic financial system.

According to the "Guidelines for the Single State Monetary Policy" in early the 1990s, averting hyperinflation by limiting extraordinarily high money growth (see Table 1) had become one of the priority objectives of the CBR's monetary policy together with other tasks such as stabilizing the financial system and the exchange rate. In the Federal Law "On the Central Bank of the Russian Federation (Bank of Russia)", which was passed in 1990, setting targets for money supply growth was indicated as one of the principal tools and methods of the Bank of Russia's monetary policy[3].

During this period the efforts to achieve macroeconomic stability were generally framed in the context of IMF-supported programs. These programs had several components (the exchange rate regime, monetary and exchange rate policies, fiscal policy and structural reforms) and implied setting intermediate targets for a number of macroeconomic (including monetary) variables regarded as nominal anchors. An underlying relationship between money growth and inflation projected in the program was a key assumption, although in practice a much more eclectic set of macroeconomic theories and modeling techniques were used to provide analytical support for the policy design (see Ghosh et al. (2005)). The CBR also studied closely the strategies of other central banks, including the monetary targeting strategy of the Deutsche Bundesbank.

The CBR's monetary policy was conducted by setting limits for the growth of the narrow monetary base[4] and other positions of the central bank's aggregated balance sheet in the Monetary Program. This included strict limits on direct loans by the CBR to the government and the commercial banks. Setting limits for money supply growth was formulated in terms of the monetary aggregate M2 (national definition) that "includes all cash and non-cash funds of resident non-financial and financial institutions (except for credit institutions), and private individuals in rubles."[5] Quarterly targets for CBR's balance-sheet indicators were set and mostly fulfilled. According to these plans, money growth was to be stabilized and subsequently slowed down. Although the CBR changed its interest rates and the reserve requirements during this period its most important tool had undoubtedly been the volume of loans provided to commercial banks and the government.

Obviously setting an adequate quantitative target for money growth was extremely complicated during the period of transition. High uncertainty and volatility of the main macroeconomic indicators caused rapid fluctuations of the demand for money. The situation was hampered even further by the lack of statistical data. Nevertheless, using elements of monetary targeting in the CBR's monetary policy helped to cope with hyperinflation, stabilized the situation in the financial sector and prevented a systemic banking crisis.

The period 1995–1998. Starting from 1995 the CBR's monetary policy framework changed considerably. Direct CBR loans to the government were discontinued. The exchange rate was

3 This clause is still present in the Federal Law "On the Central Bank of the Russian Federation (Bank of Russia)", article 35.

4 The monetary base (narrow definition) consists of the currency issued by the CBR (including cash in the vaults of credit institutions) and required reserves balances on ruble deposits with the CBR.

5 Money supply (national definition) "is defined as the sum of funds in the Russian Federation currency, intended for use as payment for goods, work and services and for the accumulation of savings by resident non-financial and financial organizations (except for credit ones) and individuals". Bulletin of Banking Statistics No 5 (216), 2011, pp. 233–234.

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used as the nominal anchor and an exchange rate band was introduced and defended by the CBR till the crisis of 1998. Domestic price stability was also mentioned as a monetary policy objective and the prevalent role of monetary expansion in determining inflation rates over the medium-term was acknowledged[6] .

The Monetary Program still included reference growth rates for the narrow monetary base, CBR's net foreign assets and net credit to the government and commercial banks, although its parameters were no longer viewed as strict targets. Under this framework, combined with the exchange rate policy, the CBR managed to bring inflation rates down to an annual 11% and money growth to 30% in 1997, although the state of the financial sector was still far from healthy, as problems with illiquidity and nonpayment of enterprises persisted, leading to widespread use of barter and monetary surrogates.

The CBR's analytical work in the area of monetary analysis in the 1990s was mainly focused on analyzing money demand, money velocity and money multiplier dynamics. Different components of money stock (including foreign-currency-denominated ones) as well as the sources of money growth were monitored. When foreign-currency-denominated deposits were legalized, in 1995 the CBR started to compile and report the dynamics of a broader monetary aggregate - broad money (or M2X)[7].

The crisis of 1998 which was due to unsustainable public finances in Russia and capital outflows from emerging countries, hit the Russian economy hard and determined the need to change the CBR's monetary policy. On the one hand, the CBR had to keep the monetary stance to prevent depreciation of the national currency and combat rising inflation. On the other hand, the dire problems in the financial sector and dysfunctions of the payment system called for liquidity-providing measures. In September 1998, the CBR abandoned the fixed-exchange-rate peg, allowed the ruble to depreciate sharply, and declared the transition to a managed floating exchange rate regime.

The period 1999–2008. In 1999 the objective of CBR's monetary policy was formulated as achieving stable economic growth in a low-inflation environment. Yet, as the capital inflows (mainly originating from the rise of oil and gas prices) increased, the CBR's commitment shifted towards exchange-rate management. Since 2003 a target for real exchange-rate appreciation was declared together with an inflation target. In 2005 the CBR introduced a bi-currency basket consisting of USD and euro (with current weights of 0.55 and 0.45 accordingly) as its operational target. In order to prevent the ruble's excessive appreciation, the CBR had to conduct substantial foreign exchange interventions which became an important liquidity-providing factor. In an environment of strong capital inflows and relatively high oil prices, the Russian economy grew strongly. From 2000 until mid-2008, the annual growth rates of M2 were above 30%.

Although the relationship between money and inflation in a relatively low inflationary environment was now less evident and the CBR no longer attempted to target money growth, the monetary aggregates retained their role as inflation risk indicators and were monitored closely. Every year the CBR published the references for M2 growth as well as the parameters of the Monetary Program in the "Guidelines for the Single State Monetary Policy". These estimates conform to the scenarios of macroeconomic development produced by the Ministry of Economy. Yet, in practice, the actual outcomes might deviate from these projections significantly. The analysis of causes and consequences of these deviations provides valuable information and is part of the analytical work in the area of monetary analysis. At this stage, the aspects of monetary analysis related to extracting information from monetary developments in order to assess the current monetary stance (as opposed to making the projections of monetary indicators contained

[6] CBR, Guidelines for the Single State Monetary Policy in 1997, p. 23.

[7] Broad money comprises cash issued by the Bank of Russia (excluding cash in vaults of the Bank of Russia and credit institutions), funds held by residents (individuals and organizations other than credit institutions) in settlement, current and deposit bank accounts denominated in rubles and foreign currencies, precious metals and all interest accrued on deposit operations.

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in the Monetary Program) started to gain importance. Naturally the relevant tools employed by the ECB for this purpose formed the basis of the analytical framework.

Money growth projections are traditionally formulated in terms of the M2 aggregate (national definition) as well as the general discussion about the monetary developments in Russia. Therefore the money demand studies conducted at the CBR originally concentrated on modeling this indicator. But as the role of monetary analysis expanded beyond the production of such projections, the need to explore the properties of other monetary aggregates and their linkages with other macroeconomic variables became apparent. In fact, the dynamics of broader aggregates that include foreign currency denominated assets are less prone to fluctuations arising from changing currency preferences and are therefore easier to interpret. Foreign currency deposits, as well as cash in foreign currency, serve as a store of value and as a safe haven during turbulent times.

The period after 2008. In recent years the CBR has adjusted the priority of its monetary policy objectives. This was partially a result of the crisis of 2008 which highlighted the impact of financial-sector imbalances on the real sector.

In 2008 the CBR declared in the "Guidelines for the Single State Monetary Policy" that lowering and subsequently maintaining low inflation is the main monetary policy objective[8]. Starting from 2009, the monetary policy horizon was extended to 3 years. The CBR also announced the gradual transition to a flexible exchange rate regime[9]. In 2010 the CBR declared that it would pay special attention to the broad analysis of money and credit developments for the purposes of financial stability and underscored the important role of credit and asset-price developments in identifying financial imbalances. In the "Guidelines for the Single State Monetary Policy in 2011 and for 2012 and 2013" it is noted that "… the Bank of Russia will pursue monetary policy by considering the situation on the financial markets and the risks arising from growth in monetary aggregates, credits and asset prices. It will pay special attention to a more comprehensive analysis of trends in monetary and credit indicators, to ensure that its timely actions in monetary policy and banking regulation and supervision help prevent imbalances in the financial sector of the economy, and thereby not only bring down inflation, but also maintain financial stability and a state of overall macroeconomic equilibrium."[10]

In the "Guidelines for the Single State Monetary Policy in 2012 and for 2013 and 2014" there is a declared intention to complete the transition to an inflation-targeting regime within a 3 year period[11]. At the same time, monetary analysis will retain its prominent role in identifying inflation risks in the medium and long-run. The CBR will also pay close attention to money, credit and asset prices developments for the purpose of maintaining financial stability[12]. As outlined by the CBR's First Deputy Chairman, Alexey V. Ulyukaev: "If you have rapid money growth you will most likely get high inflation or you could get a growth of asset prices, for example of housing or equities, that is not reflected in inflation measures …. We should cross-check inflation targeting with a monetary analysis approach. Methodologically that is what our colleagues in the ECB call two-pillars" (Ulyukaev, 2011).

Monetary analysis at the CBR therefore looks not only at price but also at financial stability, since financial imbalances have been more closely connected to high inflationary periods in Russia than in developed economies during the recent past.

8 CBR, Guidelines for the Single State Monetary Policy in 2008, I. Medium-term monetary policy principles, p. 3.

9 CBR, Guidelines for the Single State Monetary Policy in 2009 and for 2010 and 2011, I. Medium-term monetary policy principles, p. 4.

10 CBR, Guidelines for the Single State Monetary Policy in 2011 and for 2012 and 2013, I. Medium-term monetary policy principles, pp. 3-4.

11 CBR, Guidelines for the Single State Monetary Policy in 2012 and for 2013 and 2014, I. Medium-term monetary policy principles, p. 3.

12 CBR, Guidelines for the Single State Monetary Policy in 2012 and for 2013 and 2014, I. Medium-term monetary policy principles, p. 4.

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Table 1
Monetary aggregate M2 and CPI (annual growth, %)

	CPI	M 2
1992	2500	670
1993	840	410
1994	220	200
1995	130	130
1996	21,8	30,8
1997	11,0	29,8
1998	84,4	21,3
1999	36,5	57,5
2000	20,2	61,0
2001	18,6	39,9
2002	15,1	32,4
2003	12,0	50,4
2004	11,7	35,8
2005	10,9	38,5
2006	9,0	48,7
2007	11,9	43,5
2008	13,3	0,8
2009	8,8	17,7
2010	8,8	31,1

Source: CBR.

In Table 1, a comparison between inflation and annual rates of money growth suggests that the link is medium- to long-term, a short-term link being fairly difficult to establish. Empirical analyses also suggest that there should be a long-run link and that the link is closer for high-inflation regimes as discussed in Papademos and Stark (2010), chapter 1.[13] We therefore assess their co-movement for a very long time-sample and by applying filtering techniques in order to capture the trend movements and to eliminate the cyclical fluctuations. For this purpose we compile a historical dataset that although somewhat eclectic in our opinion provides an insight on inflation and money growth developments in Russia during the time span 1861–2010. This period however includes two episodes of hyperinflation: the first associated with the First World War and the Russian Revolution of 1917 and the second with the dissolution of the Soviet Union. As we do not consider these developments relevant for the objective of analyzing long-run trends in money and inflation, we deliberately remove these outliers from the data by means of the TRAMO-SEATS pre-adjustment procedure making use of a manually set sequence of deterministic variables over the periods of 1914–1923 and 1991–1993 and then apply the asymmetric Christiano-Fitzgerald filter to extract long-run trends from the data. As in Benati, 2009 we extracted the components with a frequency of oscillation over 30 years.

In Figure 1 we demonstrate the close co-movement of the two series, at the same time, the charts also suggest, however, that the strength of the correlation may be influenced by the monetary regime and the hyperinflationary regimes which – though filtered – still remain to have

[13] See, for example, Rolnick and Weber (1997) De Grauwe and Polan (2005) or Benati, 2009.

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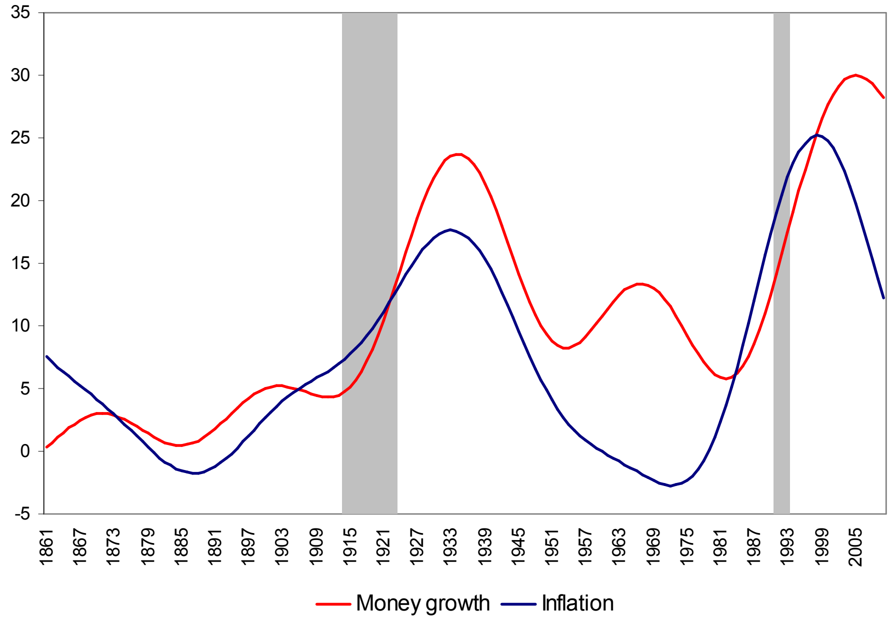
a strong influence. During the pre-soviet period the money growth and inflation rates seem to move closely. During the Soviet period of regulated prices, however, a substantial gap between money growth and inflation persisted in the 1960s and 1970s[14]. The post-Soviet period of the Russian economy was characterized by relatively high growth rates of both money and prices.

Figure 1

Long-run components of money growth and inflation, % (data over shaded periods were cleaned of outliers)



Source: authors' calculations.

3. MONETRAY DEVELOPMENTS IN RUSSIA

3.1. Types of monetary aggregates in Russia and their measurement

Definitions of monetary aggregates spread from narrow, i.e. more liquid aggregates to broader aggregates that also include less liquid components which serve the store-of-value rather than the transactional purposes of money. Moreover, definitions are influenced by the financial environment and the behavior of money holders, for example, financial institutions apart from credit institutions may also serve monetary purposes and some financial products have become so money-near that they should be included in the definition of money. While this has driven considerations for defining monetary aggregates in the euro area, broader Russian monetary aggregates reflect rather the importance of foreign-currency-denominated components.[15] M2 (national definition) is the major aggregate for the analysis and policy formulation at the CBR. Broad money (M2X), however, includes foreign-currency-denominated components (FC). This aggregate differs substantially in size and development from the aggregates that include only components denominated in national currency (NC). Over the last decades, the Russian economy has been subject to significant fluctuations in the demand for foreign currency. The flows between

[14] Interestingly, some researchers point out that the monetary overhang accumulated by the late 1980s was one of the reasons that triggered a hyperinflation spiral once prices were liberalized (see e.g. Kim, 1999).

[15] Since 2011 the CBR has published the data on deposits in national and foreign currency, set out by different sectors (financial institution (except credit organizations), public non-financial organizations, other non-financial organizations and households) in the Banking System Survey. This information provides a basis for further enhancing monetary analysis by using the data on sectoral money holdings.
See also "Sectoral structure of money holdings" (CBR, "Quarterly Inflation Review" 2011, Q1, pp. 24–26).

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ruble- and foreign-currency-denominated assets were particularly drastic during the periods of instability which impacted significantly on monetary aggregates. The recent crisis of 2008–2009 is one of the most evident illustrations and shows the need to analyze broader aggregates that partly consist of foreign currency denominated assets.

The data on the monetary aggregates M2 in national currency have been published by the CBR since 1997. The statistical sources are selected liabilities of the monthly consolidated balance sheets of Russian credit institutions and the Bank of Russia.

Two components are singled out as part of the monetary aggregate M2 (national definition)[16]:

The monetary aggregate M0 (cash in circulation) includes banknotes and coins in circulation less currency holdings (cash vaults) of the Bank of Russia and credit institutions.

Non-cash funds in national currency comprise the balances of funds kept by non-financial and financial institutions (except credit institutions) and private individuals in settlement, current, deposit and other demand accounts, including plastic-card accounts, and time-deposits opened with banks in the Russian Federation currency and accrued interest on them. Non-cash funds that are accounted for in similar accounts in credit institutions whose license has been recalled are not included in the composition of the non-cash funds.

The M1 aggregate can also be calculated from the liabilities of the consolidated balance sheet of the banking system. In our study we construct the **M1 aggregate**, which includes cash in circulation outside the banking system and transferable deposits which include current and other demand accounts (including bank card payment accounts) opened by Russian Federation residents (organizations and individuals) with the Bank of Russia and operating credit institutions in national currency[17].

Analyzing national currency monetary aggregates may be not sufficient, since financial dollarization is an important feature of the Russian economy (see Ponomarenko et al. (2013) for review). The hyperinflation that occurred in the early 1990s and the major depreciation events (most importantly, the currency crisis of 1998) increased the demand for reserve currency. Money holders however may use money for different purposes. Cash in foreign currency (mostly the USD), for example, served routinely for both transactional and store-of-value functions in the 1990s. Following macroeconomic stabilization and the increase of confidence in the banking system, the role of foreign cash has declined substantially but bank deposits denominated in foreign currency are still popular as a store of value. The shifts of currency preferences are a common reaction to exchange-rate fluctuations and increasing economic uncertainty.

The measure of money stock used by the CBR which includes foreign-currency-denominated deposits is the **broad money (M2X) aggregate.** The statistical data for this indicator was published in the Monetary Survey from 1995 to 2000 and in the Banking System Survey thereafter. Broad money comprises all the components of M2 and foreign-currency-denominated deposits.

In this study we also construct the monetary aggregate M2Y which includes foreign cash holdings in the non-financial sector. The M2Y aggregate is not published by the CBR and, as it includes cash denominated in foreign currency, the accuracy of its measurement is limited. In this study we use the indirectly measured foreign-cash holdings reported in the International Investment Position of the Russian Federation and Balance of Payments of the Russian Federation.[18] In Table 2 we summarize the components of the different monetary aggregates used in this study.

[16] Bulletin of Banking Statistics No 5 (216), 2011, pp. 233–234

[17] Data source: CBR, Banking System Survey.

[18] We use the item "Cash foreign currency/Other sectors" from the International Investment Position of the Russian Federation and the Balance of Payments of the Russian Federation.

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Table 2
Components of monetary aggregates

Liabilities		M0	M1	M2	M2X	M2Y*
Currency in circulation	X	X	X	X	X	
Demand deposits in NC		X	X	X	X	
Time and saving deposits in NC		X	X	X		
Deposits in FC			X	X		
Cash in FC				X		

* authors' definition.

In our study we also use M2X and M2Y which we adjust for valuation effects of foreign-currency-denominated components (M2X_ADJ and M2Y_ADJ). It may be sensible to do this for the purposes of monetary analysis since the fluctuations caused by the changes in the exchange rate are not linked to any real transactions and could therefore be misleading.[19] On the other hand, the wealth effect caused by these re-evaluations could still have some macroeconomic impact. We therefore analyze both types of aggregates. These were estimated as follows.

First the growth rates were adjusted:

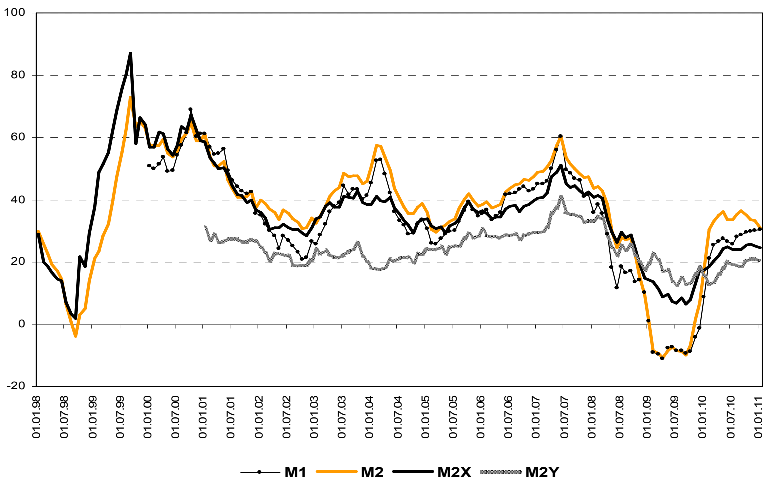
$$\Delta adj = w^{*}\Delta r + (1 - w)^{*}\, \Delta f/e \tag{1}$$

where w is the share of ruble-denominated components at the end of previous period, Δr – growth rate of ruble-denominated components, Δf – growth of foreign-currency-denominated components and e – ruble's depreciation against the bi-currency basket. The base index is then constructed using adjusted growth rates.

3.2. Evolution of different monetary aggregates and counterparts

Figures 2 and 3 show the evolution of different monetary aggregates in Russia since 1998. In Russia, distinguishing between monetary aggregates that include and those that exclude money denominated in foreign currency is particularly useful. As previously mentioned, attributing the store-of-value function mainly to deposits in foreign currency and the transactional function to foreign cash would simplify the microeconomic behavior of different money holders.

Figure 2
Monetary aggregates (y-o-y growth, %)



Source: CBR, Rosstat.

19 Russian monetary statistics so far cannot disentangle changes from transactions as it is the case for monetary data in the Eurosystem.

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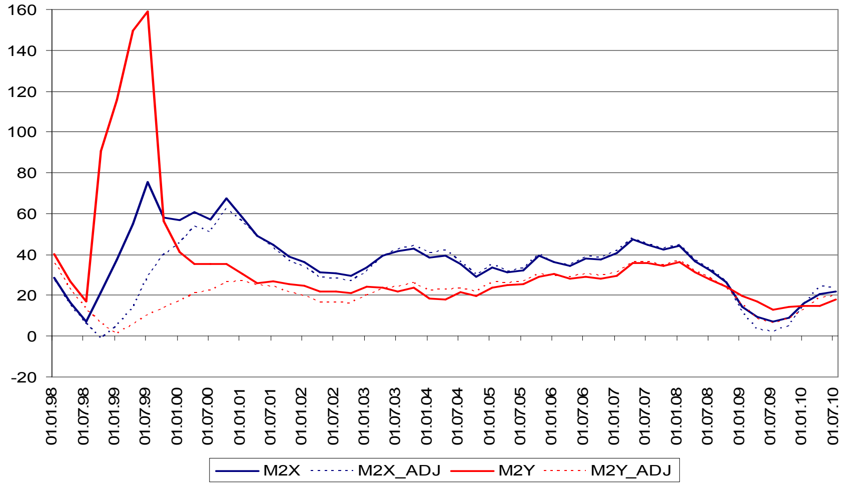
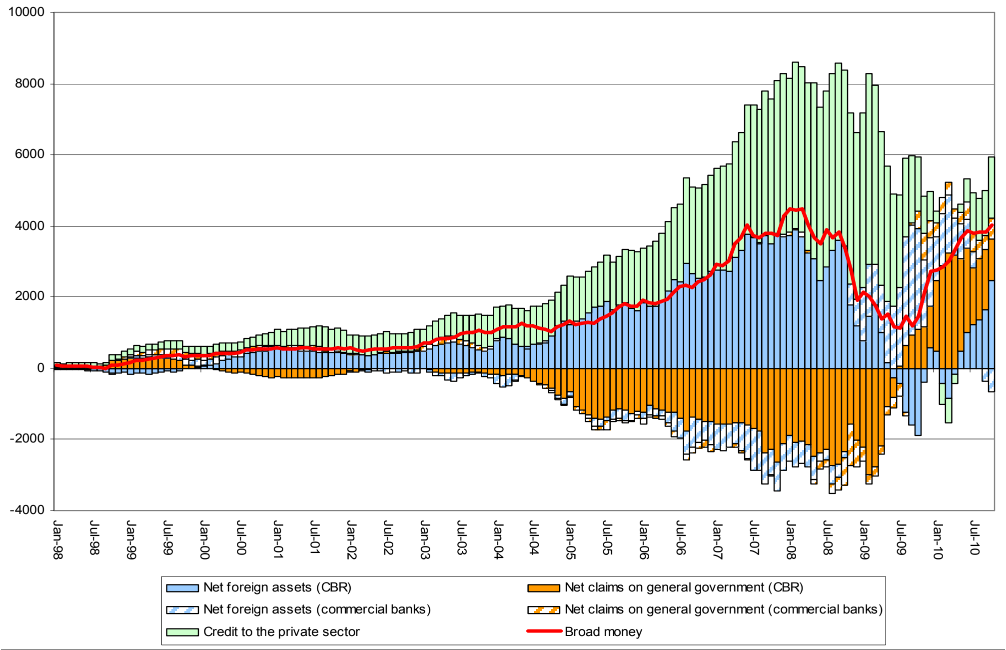
Figure 3
Headline and adjusted monetary aggregates (y-o-y growth, %)



Source: CBR, authors' calculations.

Figure 4
Money and its counterparts (annual changes, bln. rubles)



Source: CBR.

Looking at the evolution of the counterparts of Russian broad money (M2X) in Figure 4 reveals the domestic and external driving forces of monetary developments. The most important counterparts of money growth have been the CBR's foreign assets, the CBR's net claims on the government (reflecting the transactions of real domestic sector with the foreign sector) and banks' credit to the non-financial sector. Changes in the CBR's net foreign assets are generally the key driving force of changes in M2X. Changes of net claims to the general government (CBR) reflect the workings of the sovereign wealth funds, since international inflows of foreign currency are partly deposited in a sovereign wealth fund held on the CBR's balance sheet. The presence of significant exogenous growth sources means that the link between money and credit growth may not be very close – we will discuss the drivers behind different episodes of money growth later in thispaper. It also means that nominal money stock may be driven by factors totally unrelated to

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money demand fundamentals. This does not mean however that the money-demand relationship is non-existent (as money growth may trigger the adjustment of other macroeconomic variables towards new equilibrium) or that it is of no practical use. The composition of drivers behind money-stock growth indicates that the Russian economy is evidently prone to exogenous money-supply shocks (as opposed to endogenously-driven money-demand shocks). Identifying these shocks and their macroeconomic consequences is a crucial task for monetary analysis. Using money demand models to assess the degree of correspondence between realized money growth and macroeconomic fundamentals could be regarded as one of the methods of such identification.

In the early 1990s, the transformation from the planned economy in Russia was followed by galloping inflation, a deep recession, a depreciation of the national currency and large permanent government budget deficits. Money growth rates were extremely high. The new Russian banking sector at that time was just emerging and could not provide efficient financial intermediation. In these circumstances the CBR's credit to the government, to commercial banks as well as to selected non-financial enterprises was practically the only source to satisfy money demand. The direct monetization of the government budget deficit played an important role in money growth.

As the direct CBR's credit provision to the government was discontinued in 1995 the growth rates of monetary aggregates in 1996–1997 as well as inflation rates were much lower as compared to earlier 1990s.

During the 2000s, the Russian banking sector underwent a significant transformation. Although it remained small in terms of net assets to GDP, when compared to other emerging economies (Fungáčová and Solanko, 2009a), credit flows to the real sector have increased rapidly in recent years and become an important determinant of cash flows in the economy. The rapid growth of deposits (resulting in part from the cross-border cash inflows in conditions of a heavily-managed exchange-rate regime) have provided banks with a rich resource for lending. Similar conditions have been seen in Asian economies with similar monetary-policy regimes (Mohanty and Turner, 2010).

Russia turned to the fiscal mechanism of the sovereign wealth fund to absorb foreign currency from central bank interventions. In 2004 the sovereign wealth fund (the so-called Stabilization Fund which was reorganized into Reserve Fund and National Welfare Fund in 2008) was created within the Russian public finance framework. This institution proved to be very important for monetary developments and has affected the dynamics of the money stock ever since. The main source of the sovereign wealth fund's formation is taxes on oil and gas extraction and custom duties on oil exports. These funds are placed in special accounts of the Federal treasury in the CBR and are managed by the CBR. From 2005 till late 2008 the budget had a large surplus, mainly due to high oil and gas prices, which determined the accumulation of reserves on the sovereign wealth fund's accounts effectively containing money growth[20]. Changes in net foreign assets held at the CBR and net claims on the general government held at the CBR have been the driving counterparts of M2X since 1998. They reflect the functions of the sovereign wealth fund in an oil-rich economy. Its stabilizing effects, for example, are reflected in increasing positive contributions of CBR's net claims on general government after the crisis in 2008 that largely determined the recommencement of M2X growth. This reflects the buffering function of the sovereign wealth funds.

An important distinction between Russian and Asian banks was that the size of the lending booming exceeded deposit growth in 2006−2008, causing funding gaps to emerge. Russian banks relied on external borrowing to finance this gap; interbank lending in particular became dominated by transactions with foreign counterparties (Fungáčová and Solanko, 2009b).

[20] Although the CBR also used liquidity absorbing tools (such as bond issuance) the absorption through fiscal mechanisms had clearly the most important impact on the monetary stance.

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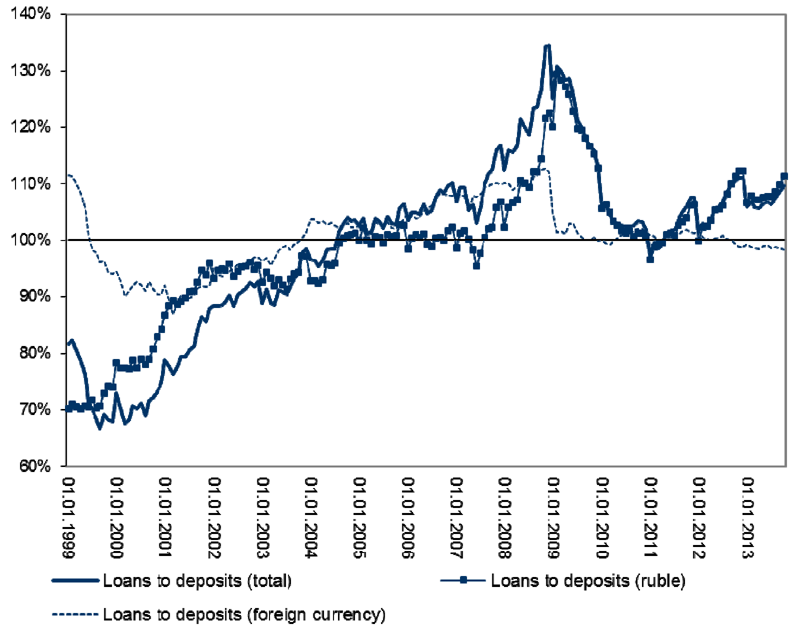
Figure 5
Loans to deposits ratio



Source: CBR.

4. MONEY DEMAND MODELS

An important aspect of the empirical properties of monetary aggregates is the existence of a stable money-demand function. The money-demand function is a fundamental relationship that captures the interactions between money and other important economic variables such as income and wealth. The role of opportunity costs is influenced *inter alia* by the depth and breadth of financial markets and the degree of substitution between domestic and foreign currencies. Thus a robust relationship between monetary aggregates and other macroeconomic variables can help to explain and interpret monetary developments. From a normative perspective, money-demand models are a starting point for developing benchmarks of the level or growth of money. In this study we are able to analyze money demand for different monetary aggregates as described in section 3.

Previous studies on money demand functions in Russia (e.g. Oomes and Ohnsorge (2005); Korhonen and Mehrotra (2010); Mehrotra and Ponomarenko (2010)) report stable money-demand relationships over the pre-crisis period. In our study we will examine if there is still a robust relationship when 2009-2010 observations are added to the sample and we will check that for different monetary aggregates. Interestingly, Oomes and Ohnsorge (2005) also conducted their estimates for several monetary aggregates and found, based on the confidence-intervals width and the recursive estimates of cointegrating vectors' coefficients that the M2Y money demand function was the most stable, while narrower ruble aggregates did not produce stable relationships. We compare these findings with more recent results.

4.1. Model specification and data issues

Our specification of the long-run real money demand in the log linear form is:

$$(m-p)_t = \beta_0 + \beta_1 y_t + \beta_2 w_t + \beta_3 OC_t + \beta_4 unc_t + \xi_t, \tag{2}$$

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where $m - p$, y and OC are the monetary aggregate deflated by the price level, the scale variable and the vector of opportunity costs accordingly. Modern money-demand studies (e.g. Greiber and Setzer (2007); Beyer (2009)) also control for the wealth effect (which as discussed in Mehrotra and Ponomarenko (2010) may be important for Russia) by adding a real-wealth variable into the money-demand function. Another addition to the traditional specification could be an uncertainty variable as in e.g. Greiber and Lemke (2005), which could also be relevant for emerging economies (see Özdemir and Saygili (2010)) particularly when attempting to model crisis developments. Recent studies by de Bondt (2009) and Seitz and von Landesberger (2010) include both wealth and uncertainty indicators into the money demand function. Therefore we add real wealth (*w*) and uncertainty (*unc*) variables into our model. We estimate four different models with real M1, M2, M2Y and $M2Y^{adj}$ as money stock variables. We do not report the results for real M2X and $M2X^{adj}$ as we fail to find any meaningful money demand relationship for these aggregates. This result is somewhat puzzling. One possible explanation is that the developments of the M2X aggregate are affected by changes of preferences between foreign-cash holdings and foreign-currency-denominated bank deposits. These changes may be difficult to model formally (at least when based only on money-demand fundamentals).

We follow Mehrotra and Ponomarenko (2010) and use a real-asset price index as a proxy for real wealth. The index is the weighted[21] average of housing and equity price indices. Housing wealth may be viewed as constituting a significant part of households' wealth. The 2002 national census found only about 3% of households rent a house or an apartment and that about 20% of households owned a secondary dwelling (mainly for seasonal use). Equities are not a significant component of household financial wealth, but their price can be viewed as a proxy for corporate wealth. As discussed in Mehrotra and Ponomarenko (2010) the rapid growth of asset prices in Russia in 2005-2007 could have positively affected transactions demand for money as transactions in asset markets increased. The increase in wealth due to the growth of asset prices may also be associated with increased demand for other liquid assets (including money) that are part of the wealth portfolio.

We have tested various indicators of uncertainty (e.g. the unemployment rate, oil-price volatility, government budget balance). Based on the models' performance and following Greiber and Lemke (2005) who propose stock-market volatility as one possible indicator of uncertainty we selected the variance of RTS index returns over rolling periods of 180 days as the metric for uncertainty. Interestingly, the interplay between this variable and various monetary aggregates may be different. Increasing uncertainty is generally associated with growing precautionary demand for money, but in case of Russia it may also result in additional demand for foreign-currency-denominated assets at the expense of ruble money stock. Therefore the positive effect on the demand for money may be more pronounced in case of broad monetary aggregates.

The choice of the opportunity-cost indicator is quite complicated in the case of Russia. The relative underdevelopment of the financial market precludes the use of money market interest rates for this purpose. On the other hand, the exchange-rate fluctuations were identified as important money-demand determinants in Russia by all previous studies as well as in other emerging market economies (see e.g. Dreger et al. (2007)). Interestingly national currency depreciation can be considered as opportunity cost only for holding ruble aggregates, since interflows between ruble and foreign-currency-denominated deposits would not affect broad money measures. In fact national currency depreciation would increase the implied ruble yield of foreigncurrency-denominated components of broad aggregates. Another opportunity-cost indicator that may be considered (as in e.g. Korhonen and Mehrotra (2010)) is the inflation rate. This leaves us with a range of variables that could be potentially used to proxy for opportunity costs/own yield.

[21] Similarly to Gerdesmeier et al. (2010) the weights are inversely proportional to the variables' volatility, i.e. Δ *Asset prices* = $\sigma_{sp}/(\sigma_{sp} + \sigma_{hp})$ Δ*Housing prices* + $\sigma_{hp}/(\sigma_{sp} + \sigma_{hp})$ Δ*Equity prices*, where σ is the standard deviation of the respective variable. The resulting weights equaled 0.86 for housing and 0.14 for equity prices and seem economically meaningful and consistent with weights used in Mehrotra and Ponomarenko (2010).

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Including all these simultaneously into the estimated relationship is hardly appropriate due to length limitations in time series. Instead we choose more parsimonious approach and construct aggregate opportunity costs/own yield measures.

The own yield of ruble components is measured by the interest rate on households' long-term ruble time deposits. The own yield of foreign currency components is the weighted average of interest rates on euro and USD deposits (with time-varying weights equal to those in the CBR's bi-currency basket[22]) plus the ruble's depreciation against the bi-currency basket over the two last quarters, which presumably proxies the exchange rate expectations. The aggregate yield of return is the weighted average (with weights proportional to the shares of ruble and foreign currency deposits in the total amount of deposits) of ruble and foreign-currency components' yields. All opportunity-cost variables are in quarterly terms.

For the money demand functions with M1 we use the aggregate yield of return as the *OC* variable and expect the β_3 coefficient to be negative, since M1 does not include appreciably remunerated components. For money-demand functions with M2 we use the exchange-rate depreciation against the bi-currency basket over the two last quarters as a proxy of the spread between ruble and foreign-currency components' yields and also expect the β_3 coefficient to be negative[23]. For money-demand functions with M2Y we use the spread between the aggregate yield of return and the realized two quarters CPI inflation rate and expect the β_3 coefficient to be positive. The overall dynamics of the resulting aggregate indicators over tranquil periods are mostly determined by changes of interest and inflation rates, but largest variations are due to exchange rate fluctuations (most notably in 1999 and 2008–2009).

We use GDP as a scale variable and the GDP deflator to calculate money and wealth variables in real terms. All variables except *OC*s and *unc* are in logs. The time sample under review is 1999Q1–2010Q2 which gives us 46 quarterly observations. The order of integration of the variables is determined based on the results of Phillips-Perron, KPSS and ADF-type test which controls for possible structural break over the crisis period (Lanne et al. (2002)) unit root tests (Table A1 in Annex A). Despite some indication from the Phillips-Perron test that $M2Y$, $M2Y^{adj}$ and y could be trend-stationary we assume that with the possible exception of *OC*s and *unc* all variables are I(1) and we therefore proceed with the cointegration analysis. This decision was supported by the test for the stationarity of the variables within cointegrated VAR conducted at later stages (Table A2 in Annex A).

4.2. Cointegration analysis

As a starting point of our analysis we refer to the most commonly applied method in testing for cointegration proposed by Johansen, 1996. The procedure efficiently includes the short-run dynamics in the estimation of the long-run model structure in the system of equations framework. We use the conventional VEC model of the form:

$$\Delta x_t = \Pi x_{t-1} + \Gamma_1 \Delta x_{t-1} + \ldots + \Gamma_p \Delta x_{t-p} + CD_t + \varepsilon_t , \qquad (3)$$

where x_t is a (K x 1) vector of endogenous variables and the Γ_p are fixed (K x K) coefficient matrices. We further assume that ε_t follows a white-noise process with $E(\varepsilon_t) = 0$. When some or

[22] While the structure of foreign-currency deposits in Russia is unavailable, other subsidiary indicators justify the use of bi-currency basket's weights for this purpose. The bi-currency basket is the operational target of the CBR and consists of the combination of USD and euro with time-varying weights.

[23] While the most obvious choice for M2 model would be to use the spread between ruble and foreign-currency components' yields this approach did not produce meaningful results (the β_3 coefficient had the "wrong" sign). The reason for that could lie with the extremely high ruble interest rates in 1999–2000 (which determined the highly positive values of the spread). Taking into account the state of financial markets and the lack of confidence in the domestic banking system at that time, these interest rates might be not fully representative as an attractive alternative to foreign currency assets. We therefore decided to disregard these interest rates. In other periods the spread was mostly determined by the exchange-rate fluctuations, as the interest rates remained stable, so there were no big differences between the two indicators.

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all of the K endogenous variables are cointegrated, the matrix Π has reduced rank r. D_t contains the deterministic terms outside the cointegrating vector, and C is the coefficient matrix associated with the deterministic terms. In our set-up, the model includes unrestricted constant and seasonal dummy variables. The lag length was set to 4[24].

Table 3
Cointegration test results: Barlett corrected trace statistic (p-value)

Model	Rank			
	0	1	2	3
M1	88.45 (0.00)	48.44 (0.04)	22.69 (0.26)	9.94 (0.29)
M2	70.81 (0.04)	43.67 (0.12)	24.67 (0.17)	6.23 (0.67)
M2Y	89.93 (0.00)	40.52 (0.20)	31.28 (0.03)	13.96 (0.08)
$M2Y^{adj}$	85.47 (0.00)	40.62 (0.20)	26.33 (0.12)	13.50 (0.10)

The tests as shown in Table 3 confirm the possibility of cointegration in all models since the rank of zero is rejected. Although there is some indication that the matrix Π may have rank 2 in the M1 model for the sake of economic interpretability we proceed by assuming 1 cointegrating relationship in all the models. The recursively-estimated eigenvalues and Hansen and Johansen (1999) fluctuations tests confirm the stability of cointegrating relationships[25] (Figures A2-A3 in Annex A). Admittedly there is considerable uncertainty regarding this specification choice that could potentially bias the model's performance as well as the results of characteristics tests. An alternative way to proceed (assuming a cointegration rank of 2) would be to identify the second cointegrating vector (such as long-run wealth growth relationship in Beyer, 2009) in addition to the money-demand relationship and examine its relevance in the comprehensive system of the simultaneous equations framework. This kind of analysis however was not undertaken in this study.

Table 4
Tests for weak exogeneity of variables: F-statistic (p-value)

Variable	Model			
	M1	M2	M2Y	$M2Y^{adj}$
M	11.50 (0.00)	4.09 (0.04)	0.15 (0.70)	0.05 (0.82)
Y	30.15 (0.00)	1.85 (0.17)	5.12 (0.02)	6.37 (0.01)
W	0.10 (0.76)	15.04 (0.00)	8.16 (0.00)	9.72 (0.00)
OC	0.09 (0.77)	0.73 (0.39)	63.15 (0.00)	53.43 (0.00)
UNC	0.40 (0.53)	3.16 (0.08)	0.06 (0.81)	0.37 (0.54)

Null hypothesis: variable is weakly exogenous.

24 Most of the traditional information criteria would indicate that a longer lag-length is preferable. But for the reasons of parsimony given the short time sample and given the quarterly data used we limit the lag length to four. Later we examine to what extent the lag-length choice influences the cointegrating vectors.

25 At this stage we concentrate on the analysis of long-run relationship and therefore excluded the short-run part from the stability tests. The performance of short-run money demand models are discussed elsewhere in this paper.

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Although the analysis of the dynamic relationship between money and other macroeconomic variables is beyond the scope this paper we do examine the weak exogeneity tests based on the VEC model reviewed and show the test results in Table 4. There are notable differences in the results for different models: while the weak exogeneity of narrower ruble aggregates is rejected, the developments in the broader aggregates seem to be unaffected by the adjustment resulting from the cointegration relationship. This result may contradict the conventional theory associated with the money-multiplier concept which would presume narrow aggregates to be exogenous and broader ones to be endogenous. Yet these findings may be in line with the peculiarities of money-supply factors in Russia. We will further discuss the performance of the models in explaining money-stock developments later in this paper.

Instead of affecting money, the adjustment occurs through other variables such as GDP or real wealth. The results for OC variables are mixed – they seem to be weakly exogenous in the M1 and M2 models and endogenous in M2Y and $M2Y^{adj}$ models.

The cointegration vectors are estimated by the simple two-step estimator (S2S). As Brüggemann and Lütkepohl (2005) show, this estimator produces relatively robust estimates in short samples. The lag length is set to 4. Most of the cointegrating vectors estimated using different lag lengths were relatively robust.[26]

We cross-check the results obtained with S2S method by estimating the cointegration vectors using Fully Modified-OLS (Philips and Hansen (1990)) in a parsimonious single equation set-up. We use pre-whitening with the lag length determined by Schwarz criteria and Barlett kernel with the cut-off determined by the automatic Andrews (1991) procedure.

The cointegration vectors are estimated in the presence of unrestricted constant and seasonal dummy variables. The results are shown in Table 5.

Table 5
Cointegration vectors (t-statistics)

Variable	Estimation Method	Model			
		M1	M2	M2Y	M2Yadj
M	S2S	1	1	1	1
	FM-OLS	1	1	1	1
Y	S2S	-1.65 (-37.3)	-2.38 (-20.2)	-0.38 (-4.09)	-0.63 (-12.3)
	FM-OLS	-1.76 (-12.6)	-2.6 (-13.1)	-0.61 (-1.68)	-1.05 (-4.85)
W	S2S	-0.47 (-13.8)	-0.34 (-3.49)	-0.88 (-11.2)	-0.67 (-15.5)
	FM-OLS	-0.48 (-4.45)	-0.29 (-1.68)	-0.54 (-1.81)	-0.23 (-1.31)
OC	S2S	2.07 (13.1)	3.73 (9.15)	-3.47 (-5.34)	-1.62 (-4.89)
	FM-OLS	0.93 (3.17)	0.84 (1.25)	2.18 (2.13)	-0.4 (-0.72)
UNC	S2S	-118 (-8.25)	-45.1 (-1.81)	-67.4 (-2.13)	-128.5 (-7.61)
	FM-OLS	-8.4 (0.48)	-44.7 (-1.21)	-152 (-3.16)	-109.8 (-3.86)

[26] The results are presented in the working paper (ECB WP 1471) version of this paper.

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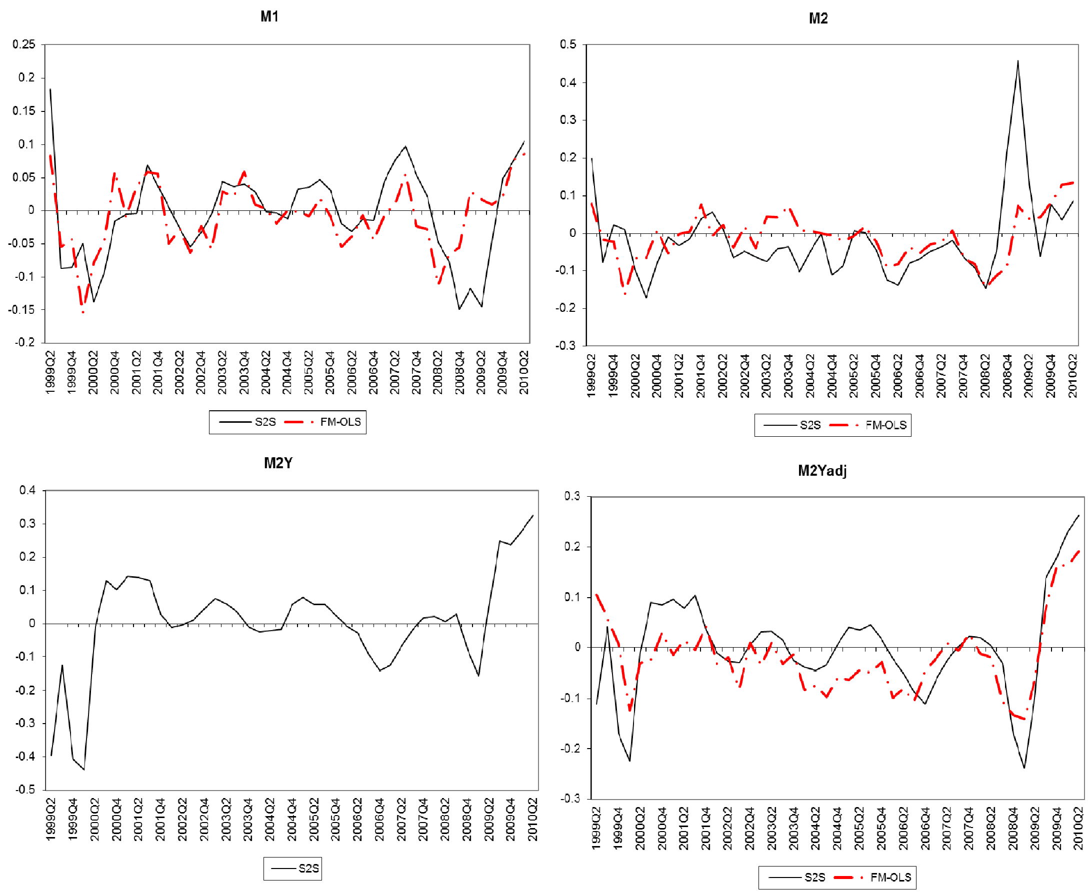
The parameters estimated with S2S method are statistically highly significant and economically meaningful. The growth of GDP and real wealth increases money demand. Interestingly, there are striking differences in income elasticities between the models. The M1 and M2 models retain the feature of high income elasticity, which was also reported in all the previous money demand studies for Russia. This peculiarity is usually associated with ongoing institutional changes such as financial deepening and the return of confidence in the national currency. In the cases of M2Y and M2Y^{adj} however the income elasticities are lower, while the parameters for wealth are somewhat higher. The sum of the income and wealth parameters is only slightly higher than unity[27]. In fact these results are consistent with the parameters reported in Greiber and Setzer (2007) for the euro area and the US and in Seitz and von Landesberger (2010) for the euro area. These results seem to be thought-provoking as they show how differently monetary developments in Russia could be interpreted when different money stock measures are used. The opportunity-cost variables all have the expected signs. The increase of uncertainty has a positive effect on money demand. As could be expected it seems to be less evident in case of ruble aggregates.

These findings are generally confirmed by FM-OLS estimates for M1, M2 and M2Y^{adj} models, although the uncertainty variable in the M1 model and OC variables in M2 and M2Y^{adj} models are statistically insignificant. Nevertheless we proceed with further analysis of this cointegration vectors as they are economically meaningful. We exclude the FM-OLS cointegration vector for the M2Y model that displays the "wrong-signed" OC variable coefficient.

Figure 6
Monetary overhangs[28]



27 The sum of coefficients equals 1.26 and 1.3. Interestingly, Oomes and Ohnsorge (2005) report an income elasticity of 1.2 for M2Y money demand function without wealth.

28 The monetary overhangs were computed (using seasonally adjusted data) as demeaned error correction terms from the estimated cointegration relationships.

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In order to test the robustness of the results we estimate the cointegration vectors recursively to check if the point estimates remain stable as the post-crisis observations are added into the sample (Figures A3–A6 in Annex A). The recursive estimates of income and wealth elasticities are relatively stable in all models irrespectively of the estimation method (with the exception of income elasticity in S2S M2Y model which was insignificant if estimated using only pre-crisis sample). The OC and uncertainty recursive coefficients displayed considerable fluctuations but still seemed meaningful in the models for ruble M1 and M2 aggregates. The recursive estimates of OC variable coefficient in M2Y and M2Y^{adj} models reveal, however, that the OC variables only started to enter the cointegration relationship with the "right" sign after the large number of post-crisis observations had been added to the estimation sample. This result may indicate that the relationship between broader monetary aggregates and OCs is more complex than implied by this money demand relationship or that the financial returns indicators do not fully represent the OCs in the Russian economy. On the other hand, given the limited variation of OCs before the crisis and relatively short time sample we can not rule out the possibility that adding the observations characterizing the opposite phase of the economic cycle was just necessary to disentangle the true effect of OCs on money demand.

We may examine monetary overhangs derived from the cointegrating relationships as the measures of excess liquidity and as shown in Figure 6. The choice of cointegration vector's estimation method does not seriously change the outcome here. With the exception of fluctuations in the beginning of the sample and the hikes of S2S M2 overhang in early 1999 and 2009 (determined by the sharp exchange rate depreciation episodes) the dynamics of the overhangsseem meaningful. They fluctuate evenly around zero, pick up in 2006 before plummeting to some very low levels in 2008–2009. Then, as money growth picked up while money demand fundamentals' (particularly real asset prices) remained weak, the monetary overhangs climbed to unprecedentedly high levels, in particular for M2Y and adjusted M2Y.

4.3. Error correction models

The short-run money-demand models are formulated as conventional ECMs of the form:

$$\Delta(m-p)_t = \alpha_0 + \alpha_1 ec_{t-1} + \sum_{j=1}^{2} \alpha_{2j}\Delta(m-p)_{t-j} + \sum_{j=1}^{2} \alpha_{3j}\Delta y_{t-j} +$$

$$+ \sum_{j=1}^{2}\alpha_{4j}\Delta w_{t-j} + \sum_{j=1}^{2} \alpha_{5j}\Delta OC_{t-j} + \sum_{j=1}^{2} \alpha_{6j}\Delta unc_{t-1} + \sum_{i=1}^{3}\sigma_i D_{it} + \varepsilon_t$$

where *ec* is the error correction term and D_i are the seasonal dummy variables. The equations include two lags of real money growth. The short-run part of the equations also contains up to two lags of first differences of other explanatory variables (these are eliminated if the respective t-statistics are smaller than 1.67). Conventional tests do not find serial correlation or ARCH effect in the equations' residuals. The α_1 coefficients as given in Table 6 are of most interest as they show that real money growth adjusts in accordance with the cointegrating relationship.

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Table 6
ECMs' α_1 loading coefficients (t-statistics)

Estimation period	Cointegration vector	Model			
		M1	**M2**	**M2Y**	**M2Y^{adj}**
1999Q1–2008Q2	S2S	-0.39 (-3.84)	-0.05 (-0.88)	0.05 (1.59)	-0.03 (-0.44)
	FM-OLS	-0.47 (-3.61)	-0.23 (-2.11)	-	-0.4 (-4.00)
1999Q1–2010Q2	S2S	-0.24 (-2.82)	0.01 (0.11)	0.07 (2.25)	-0.02 (-0.56)
	FM-OLS	-0.28 (-2.37)	-0.03 (-0.35)	-	-0.24 (-3.17)
1999Q1–2010Q2 (with dummy variables)	S2S	-0.31 (-3.70)	-0.05 (-0.98)	0.06 (1.93)	-0.01 (-0.16)
	FM-OLS	-0.41 (-3.44)	-0.19 (-2.19)	-	-0.2 (-2.65)

At first we estimate the ECMs on the pre-crisis period prior to 2008Q3. The loading coefficient in the M1 and M2 models is large and statistically highly significant (although the FM-OLS cointegrating vector is clearly more relevant for short-run M2 developments than S2S estimates). Quandt-Andrews breakpoint tests indicate that the models are stable over this sample. When the post-crisis observations are added to the time sample, the loading coefficients deteriorate notably (although in case of M1 it is still significant). The recursive estimates of loading coefficients show that their instability coincided with crisis developments (Figure A7 in Annex A). We therefore also examine the ECMs' estimates with the period of 2008Q3–2009Q1 covered with dummy variables. Under this set-up the estimates of the loading coefficient do not change significantly in comparison with the pre-crisis sample estimates.

This result could be expected, given the drastic and unpredictable fluctuations of money stock during the most severe phase of the crisis. The rapid return of dollarization, for example, could not be captured by the exchange rate variable since ruble's depreciation expectations were much stronger than implied by the gradual CBR-controlled depreciation rates. If we assume that the model's error during the crisis was due to the error in measuring exchange-rate expectations we may illustrate this by solving the model for M2[29] back and finding the exchange-rate variable value that implied no error in the model's estimate of money stock. Over most of the sample this estimate would have no economic meaning. Yet, during the depreciation episode this estimated exchange rate variable's value could be used to assess these unobserved expectations.

[29] We used the FM-OLS model for M2 estimated over the 1999Q1–2008Q2 time sample.

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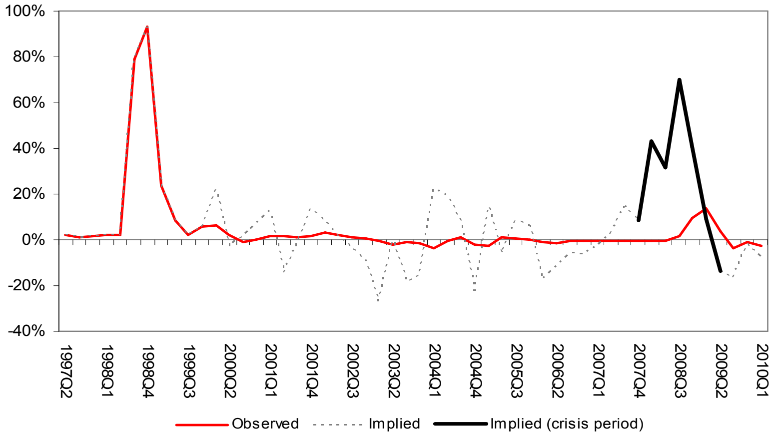
Figure 7
Exchange rate expectations: observed proxy (the average quarterly ruble's depreciation over the last two quarters) and the estimate implied by the model



Source: authors' calculations.

The results of this exercise, as shown in Figure 7, indicate that the expected ruble's depreciation which would be consistent with the intensity of dollarization was higher than the one actually realized. In fact the market participants seemed to expect a depreciation similar to the one that took place during the previous crisis of 1998.

The results of ECMs' estimation for broader aggregates are more ambiguous. In contrast to ruble aggregates the broader M2Y and M2Y^{adj} seem to be unaffected by the cointegrating relationship (at least by those estimated with the S2S method). The α_1 estimates are statistically insignificant under any set-up and in the case of M2Y the loading coefficient is positive which is clearly implausible. We believe that this difference arises from the fact that the nominal volumes of ruble aggregates are quite sensitive to changes of transactional needs and opportunity-cost fluctuations (as households are eager to switch between currencies or between cash and bank deposits). There are however fewer means for nominal volumes of broad aggregates to adjust as their dynamics is only partially determined by demand-driven processes (i.e. financial intermediation) and there are virtually no assets outside M2Y aggregate that are widely used for savings purposes. There still is the chance that the real money stock would adjust due to the increase of the price level, but given the relatively short period under review and the scope of the nominal money-supply shocks whicht took place during this period, it is unsurprising that such adjustment could go unnoticed by the econometric model. Yet, we can not rule out the possibility that broad aggregates may be driven by money demand completely as the ECM based on FM-OLS cointegration vector estimate for M2Y^{adj} performs satisfactorily and is not drastically affected by the crisis.

We can summarize our findings as follows. The long-run money demand relationship may be established for M1, M2, M2Y and M2Y^{adj} aggregates. The parameterization of these relationships is notably different, although it is impossible to discriminate between them from a theoretical viewpoint since all sets of parameters might be plausible under certain assumptions. Contrarily to Oomes and Ohnsorge (2005), the narrowest M1 aggregate performs at least as good as the broader aggregates. In fact the recursive estimates of the cointegration vector of M1 money-demand relationship seems to be more stable than those estimated for broader monetary aggregates, in which cases the robustness is questionable. The short-run model of the demand for M1 is obviously the best performing, while M2Y developments seem to be ambiguously affected by the money-demand relationship. Although given the exogenous nature of the sources of nominal money growth in Russia and the underdevelopment of the alternative financial assets that could be used for savings purposes beyond those included into M2Y, this last finding seems plausible.

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5. THE INFORMATION CONTENT OF MONEY FOR INFLATION AND FOR REAL ECONOMIC DEVELOPMENTS

As described in Papademos and Stark (2010), chapter 4, extracting the signals in monetary developments is an essential piece of information for policy-makers, though it may be difficult at times because of the short-term volatility of monetary aggregates. A key issue for central banks aiming to maintain price stability is the information content of money for future inflation. Understanding the behaviour of loans and monetary aggregates during the business cycles, however, can add useful insights for analyzing the driving forces of real economic activity. In this part we present two models that shed some light on these questions in the case of Russia.

5.1. Money-based inflation risk indicators

We use money-based inflation-risk indicators in order to assess the long-run link between monetary growth and inflation. Thus, the assessment of risks to price stability should rather focus on the persistent or lower frequency movements of the monetary developments, or – to repeat the wording of the ECB here – on the "underlying monetary growth". As outlined in Papademos and Stark (2010, p. 209) and as recognized by the ECB's two-pillar strategy, it is important to stress that money alone is insufficient for collecting information on future price developments. Money, nevertheless, is a necessary component and a starting point is to use a reduced-form inflation-forecast equation in the spirit of Stock and Watson (1999). As applied by Nicoletti-Altimari, 2001 or Hofmann, 2008 and described in Fischer et al.(2008) we use an augmented autoregressive equation for forecasting consumer price inflation, π, at time t+h with the information embodied in monetary indicators x at point t:

$$\pi^{k}_{t+h} = \beta_0 + \beta_1(L)\pi_t + \beta_2(L)x_t + \varepsilon_{t+h} \tag{5}$$

In the case of Russia and similarly to Papademos and Stark (2010, p. 220, 221), we use quarterly data and capture the idea of "underlying money growth" with two smoothing measures.

Firstly, the key explanatory variable is the (weighted) average of monetary developments over several periods as implied by β_2(L). The number of lags, just as in the case of β_1(L), are selected using the Schwartz information criterion (with a maximum of 3).[30]

Secondly, k denotes the number of quarters over which the inflation term is calculated in annualised terms. A specification with k = 4 would generate point forecasts for annual CPI inflation at any specific horizon chosen for h, while k > 4 represents annualised inflation over a period longer than one year, and thus, can be seen as a smoothing measure. In this study k ranges between 4 and 6 for the chosen six-quarter-ahead forecasts in order to produce indicators that are informative about the medium-term trends in inflation.

These trends will influence developments at horizons well beyond 12 quarters ahead, thus indicating a time horizon, at which inflation ultimately becomes a monetary phenomenon according to the ECB's monetary analysis.

We also test the performance of alternative explanatory variables under the same model set-up. We use a number of readily available variables from the categories of economic indicators that are often regarded as important inflation determinants in Russia (see e.g. Oomes and Dynnikova (2006)): quarterly GDP growth, GDP gap (recursively calculated with Hodrick-Prescott filter, λ = 1600), quarterly growth of nominal effective exchange rate (NEER) and quarterly growth of ruble oil prices. Additionally we test the performance of tri-variate economic indicators model

[30] The selection of the lags in a purely data-dependent manner implies that the precise specification of a model can change from one forecast to the next. The disadvantage of more difficult comparability between forecasts is accepted in order to have a less complex model.

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including both GDP and NEER growth as regressors that may be regarded as a proxy for a Phillips curve relationship.

Table 7 presents the results of the accuracy of the forecasts for the model based on different monetary aggregates, their median forecast, economic indicators models and the pure autoregressive model for inflation (AR). It is expressed as the ratio between the Root Mean Squared Error for the money-based forecasting model (RMSE) and the Root Mean Squared Error of the Random Walk (RMSE_RW) for different monetary aggregates in Russia. As suggested in Hofmann, 2008 we also report the results of Diebold and Mariano (1995) test (with HAC correction) on the equality of the forecast errors.

Table 7
RMSE/ RMSE_RW (p-values of the tests of the hypotheses that RMSE = RMSE_RW in parentheses)

	k = 4	k = 5	k = 6
M1	**0.51**	**0.42**	**0.3**
	-0.1	*-0.04*	*-0.01*
M2	**0.88**	**0.59**	**0.45**
	-0.62	*-0.02*	*-0.01*
M2X	**0.37**	**0.36**	**0.51**
	-0.03	*-0.02*	-0.01
M2X adjusted	**0.59**	**0.44**	**0.38**
	-0.06	*-0.02*	*-0.01*
M2Y	**0.96**	**0.84**	**0.67**
	-0.3	*-0.01*	*-0.05*
M2Y adjusted	**0.84**	**0.89**	**0.77**
	-0.03	*-0.39*	*-0.02*
Median forecast of monetary models	**0.4**	**0.4**	**0.44**
	-0.01	*-0.02*	*-0.01*
Δ GDP	**1.17**	**0.83**	**0.69**
	-0.57	*-0.05*	-0.05
GDP gap	**1.04**	**1.05**	**0.87**
	-0.59	*-0.6*	*0*
Δ NEER	**0.79**	**0.82**	**0.74**
	-0.23	*-0.09*	*-0.02*
Δ GDP, Δ NEER	**0.83**	**0.67**	**0.72**
	-0.2	*-0.1*	*-0.05*
Δ Oil prices	**0.98**	**0.89**	**0.92**
	-0.62	*-0.01*	*-0.27*
AR	**0.82**	**0.69**	**0.62**
	-0.07	*-0.02*	*-0.02*

Based on this measure most of the money-augmented inflation forecasts improve the accuracy compared to the Random Walk and autoregressive model and also outperform the economic indicators. The most informative monetary aggregates are M1 and M2X (both adjusted and

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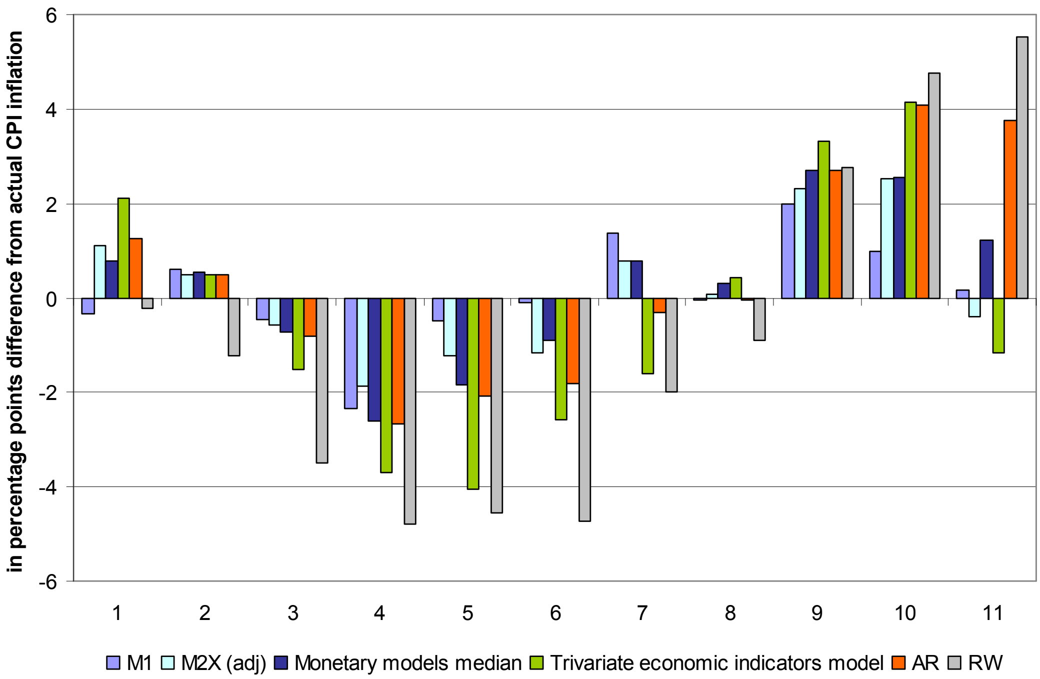
unadjusted). In general, monetary models seem to perform better when forecasting on longer horizons.

Figure 8 shows the errors of the six-quarter-ahead forecast for k = 6 compared with the actual CPI inflation outcomes for selected approaches.

Figure 8
Forecast Errors of Inflation Indicators for Russia (2007Q4–2010Q2)



1 – 2006Q2–2007Q4
2 – 2006Q3–2008Q1
3 – 2006Q4–2008Q2
4 – 2007Q1–2008Q3
5 – 2007Q2–2008Q4
6 – 2007Q3–2009Q1
7 – 2007Q4–2009Q2
8 – 2008Q1–2009Q3
9 – 2008Q2–2009Q4
10 – 2008Q3–2010Q1
11 – 2008Q4–2010Q2

Source: authors' calculations.

The forecast errors of the money-augmented models are relatively small compared to the alternative models for most periods. The fact that monetary models performed relatively well may reflect certain complications in forecasting inflation in Russia with traditional economic variables. For instance while oil-price growth is cost increasing factor[31] it is also an important income factor and is usually associated with reduced inflation due to ruble's appreciation. Extracting relevant information from (quite volatile) output growth rates in conditions of ongoing transformation of the economy (due to both the emerging market nature of the Russian economy and its susceptibility to external shocks) may also be challenging. In this context, money-based inflation-risks indicators may prove particularly useful. It should, however, be kept in mind that the forecasting exercise is restricted to a fairly short and special time window due to the availability of data.

5.2. The informational content of money for real economic developments

Using money as an information variable for real sector developments is another important aspect of analyzing monetary aggregates. In Brand et al. (2003) narrow money, here M1, seems crucial for cyclical developments in the euro area compared to other macroeconomic and financial

[31] Interestingly, the link between world oil prices and (monopolized and regulated) domestic fuel market's prices may be less evident in Russia.

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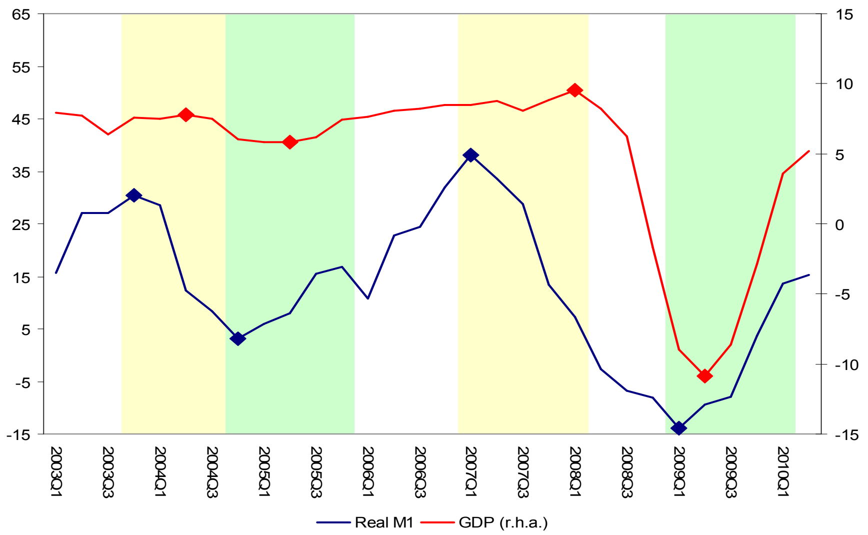
variables commonly used for forecasting real economic activity. Interestingly, this is in contrast to most studies conducted for the U.S. Thus, we investigate this question for Russia. The robust relationship between money and real growth may prove valuable for the purposes of assessing the general trends of the real sector developments as well as for more specific short-term forecasting. Indeed, some central banks underscore the importance of the close link between money and actual spending in their routine analytical work (see e.g. Bank of England, 2008 or ECB, 2008).

The exploration of the link between money and credit aggregates and economic activity has also intensified when attempting to explain drastic GDP fluctuations in Russia. In fact the analysis of the linkage between the turning points in output and real money stock dynamics conducted at the ECB in 2008 seems to be perfectly applicable for Russia. As shown on Figure 9, over the last decade every turning point in the annual growth rates of real M1 was followed by the turning point in the annual growth rates of real GDP within the next 4 quarters.

Figure 9
Real M1 and real GDP annual growth rates (%) and turning points[32]



Source: CBR, Rosstat.

This type of analysis however makes use of only several observed episodes. We will also examine to what extent the monetary aggregates can be used to forecast the real sector developments in a more linear fashion.

5.2.1. Model specification and data issues

Our empirical approach is closely related to Brand et al. (2003). We formulate the regression of the form:

$$\Delta y_t = \alpha_0 + \alpha_1 \Delta_4 m_{t-1} + \alpha_2 oil_t + \alpha_3 \Delta_4 reer_t + \alpha_4 \Delta y_{t-1} + \alpha_5 \Delta y_{t-2} + \varepsilon_t, \qquad (6)$$

where Δy is the change of seasonally adjusted GDP and $\Delta_4 m$ is the 4-quarters change of the real monetary aggregate (deflated with GDP deflator). As the model with M2Y has a "wrong" (negative) α_1 coefficient we exclude it from further analysis. We add real ruble oil prices (*oil*) and the 4 quarters changes of the real effective exchange rate (*reer*) as control variables[33] that

[32] The turning points were identified as local maxima/minima on the rolling-window periods of 9 quarters.

[33] The conventional choice of the control variable for this regression are yield spreads. Adding the spread between interest rates on loans and deposits and real interest rate on loans produces counterintuitive results (positive coefficients) and we therefore removed them.

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are commonly used to explain GDP fluctuations in Russia (see e.g. Rautava, 2009). The lagged changes of GDP prevent the autocorrelation in the residuals. All variables are in logs. The time sample was set to 1999Q1–2006Q4. The Phillips-Perron and KPSS unit root tests yield conflicting results, as the Phillips-Perron test suggests that annual changes of monetary aggregates may have unit roots (Table B1 in Annex B). However, based on the visual observation of the time series and given that there is no possibility of spurious regression since the left-hand-side variable is clearly stationary, we proceed by estimating the regression.

5.2.2. Results

Firstly we examine the in-sample performance of the models. For this purpose we compute one-step-ahead static forecasts. Following Fischer et al. (2008) we use the mean squared forecast error (MSFE) as the criteria. We also formulate the benchmark models to assess the usefulness of taking into account the monetary aggregates. As in Fischer et al. (2008) these are the simple autoregressive models (AR) and the naïve random walk forecast (RW). The third benchmark model (X) equals the model as stated in equation (3), but excludes the monetary indicator.

Table 8
In-sample forecasting performance

	MSFE	$MSFE/MSFE^{AR}$
AR	0.00007713	1.00
RW	0.00010732	1.39
X	0.00003282	0.43
M1	0.00003143	0.41
M2	0.00002816	0.37
M2X	0.00003210	0.42
M2X_adj	0.00003103	0.40
M2Y_adj	0.00003190	0.41

On the in-sample period the models with money are performing better than naïve models, although their MSFE is only marginally smaller than MSFE of the X-model that does not include money as shown in Table 8. In fact the statistical significance of money in these equations is rather low, with only M2 being significant at the 5% level. We then proceed by making the out-of-sample static forecasts over the period of 2007Q1-2010Q2. As in Fischer et al. (2008) we report both components that determine the MSFE, the standard deviation of the forecast and the forecast bias.

Table 9
Out-of-sample forecasting performance

	MSFE	$MSFE/ MSFE^{AR}$	Bias	Standard deviation
AR	0.00057	**1.00**	-0.011	0.005
RW	0.00064	**1.11**	0.005	0.020
X	0.00072	**1.26**	-0.016	0.009
M1	0.00054	**0.95**	-0.014	0.010
M2	0.00050	**0.88**	-0.015	0.012
M2X	0.00066	**1.16**	-0.016	0.010
M2X_adj	0.00076	**1.33**	-0.017	0.012
M2Y_adj	0.00082	**1.44**	-0.018	0.011

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Over the out-of-sample period, only M1 and M2 models are performing better than naïve benchmarks. Yet, the models that include M1, M2 and M2X perform notably better than the X model which does not include money. These results should be interpreted with caution as the time sample analyzed includes the crisis developments and may therefore not be fully representative. Table 10 shows the results when excluding the period of 2008Q3–2009Q2 (i.e. the period of recession) from the sample.

Table 10
Out-of-sample forecasting performance (excluding recession)

	MSFE	MSFE/ MSFE^{AR}	Bias	Standard deviation
AR	0.00005	**1.00**	-0.001	0.002
RW	0.00038	**7.21**	0.016	0.012
X	0.00009	**1.75**	-0.004	0.006
M1	0.00007	**1.34**	-0.004	0.006
M2	0.00010	**1.92**	-0.006	0.006
M2X	0.00009	**1.66**	-0.005	0.006
M2X_adj	0.00012	**2.28**	-0.006	0.008
M2Y_adj	0.00013	**2.49**	-0.006	0.008

The results of the out-of-sample forecast over the relatively calmer periods indicate that the models' performance are not impressive as none of them was able to outperform the autoregressive model. The M1 and M2X models are still forecasting better than the X model, meaning that money developments contain some useful information beyond that contained in the control variables.

Summarizing, we find only limited usefulness of using monetary aggregates to forecast GDP growth over the pre-crisis period in the linear model framework. The performance of the estimated models generally deteriorates during the crisis and post-crisis periods and is worse compared to naive models. Yet, the models which contain money seem to perform better than the one that does not. This is particularly so when the periods of recession are included, although to what extent this finding is representative remains unclear. As could be expected, the best-performing monetary aggregate appears to be M1.

6. CONCLUSIONS

Tools and techniques of the ECB's monetary analysis can give valuable input to the conduct of monetary policy at other central banks, if institutional, economic and financial differences are taken into account. We take the case of the Bank of Russia and analyze the changing role of money in its monetary policy.

In the core part of our paper we derive stable money demand functions that are related to income and wealth and to a lesser extent to opportunity costs and uncertainty. Estimations of narrower aggregates that only include components denominated in national currency seem to be more stable than broader aggregates. It signals that monetary developments are influenced by factors that go beyond the usual money-demand factors, such as the buffering function of the sovereign wealth fund in case of Russia. This makes the interpretation of monetary overhangs and the policy implications that can be drawn from them more complex since the impact of the sovereign wealth fund on monetary development is already a policy reaction. Eventually, it should be kept in mind that the concept of monetary overhangs are a starting point for an analysis that also focuses on changes in the stocks rather than analyses that solely focus on changes in

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the flows. Additionally, we show how money demand functions can be used to derive implied exchange-rate depreciation expectations as compared to actual exchange-rate depreciation.

In the last section, we present results that deliver some information content of money for inflation and real economic development. As in case of the ECB's regular monetary assessment we measure money-based inflation risk indicators and compare the performance of different monetary aggregates with naïve and univariate inflation models as well as inflation models with alternative economic variables. The results are promising, though we leave it for future analysis to assess their performance over time. The results of the information content of money for real economic developments is fairly limited, however, in line with results for the euro area, the narrow monetary aggregate seem to perform relatively better compared to broader aggregates. This, however, should not be seen as a negative feedback to the ECB style monetary analysis since monetary aggregates are rather used as an indicator for the turning points in real GDP at the ECB. During our sample period this purposes is actually served fairly well for Russia.

We conclude that the ECB style monetary analysis gives valuable input to the analysis at the CBR. Monetary analysis, however, is an evolutionary process, so within an economy over time as well as across economies changes of the economic and financial environment have an impact on the analysis and on the policy conclusions that can be drawn from it. The case of Russia furthermore highlights that monetary analysis should not be minimized to a purely technical exercise, but that it needs and enforces institutional knowledge of the financial sector.

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ANNEX A

Figure A1
Variables used in the model

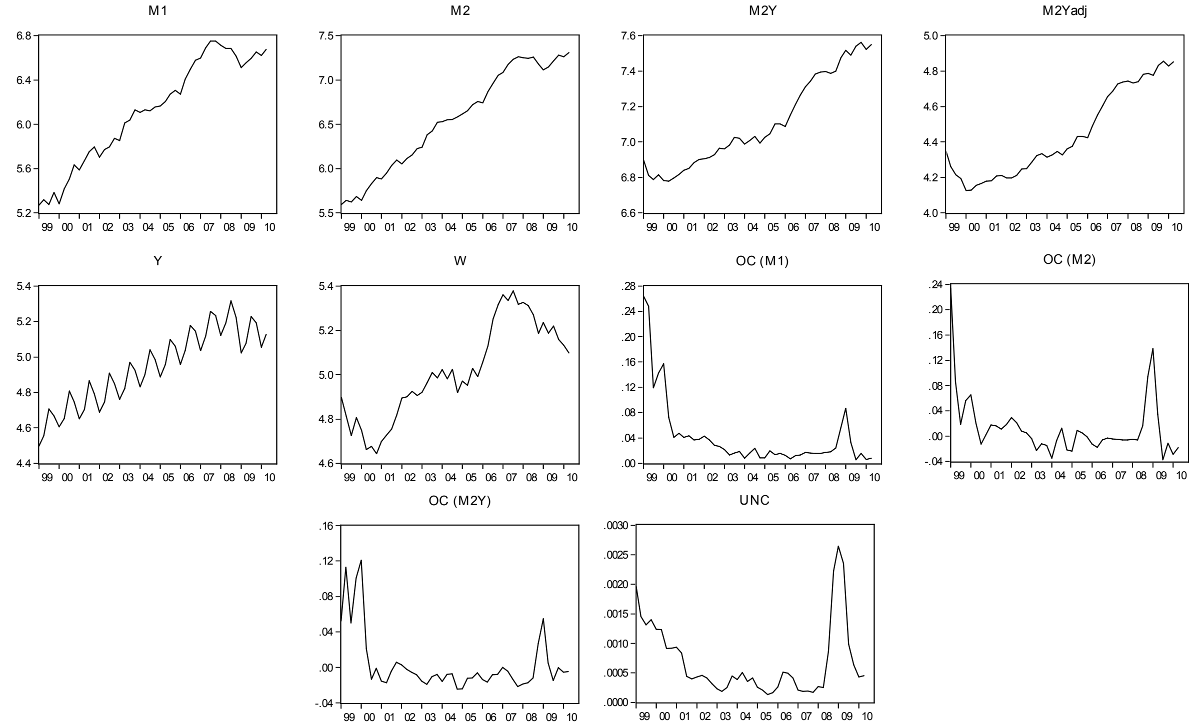

Table A1
Results of the unit root tests (for PP and KPSS tests the bandwidth is determined by automatic Andrews, 1991 procedure; the unit root test with structural break is conducted in the presence of seasonal time dummy variables and the shift type dummy variable in 2008Q4 (the impulse type dummy variable when the variables are in differences) with lag length set to 4)

Variable	**Test specification**	**PP test statistic (p-value)** Null hypothesis: variable has unit root	**KPSS test statistic** Null hypothesis: variable is stationary	**ADF type unit root test with structural break (Lanne et al. (2002)) test statistic** Null hypothesis: variable has unit root
M1	Levels (constant)	-0.68 (0.84)	0.45^{*}	-1.78
	Levels (constant and trend)	-1.80 (0.69)	0.16^{**}	-2.73
	1st differences (constant)	-7.47 (0.00)	0.14	-3.64^{**}
M2	Levels (constant)	-0.49 (0.88)	0.66^{**}	-1.4
	Levels (constant and trend)	-1.78 (0.70)	0.17^{**}	-2.15
	1st differences (constant)	-5.88 (0.00)	0.14	-3.06^{**}
M2Y	Levels (constant)	1.00 (0.99)	0.67^{**}	-0.22
	Levels (constant and trend)	-3.95 (0.02)	0.15^{**}	-0.77
	1st differences (constant)	-5.16 (0.00)	0.49^{**}	-2.9^{**}

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Variable	Test specification	**PP test statistic (p-value)** Null hypothesis: variable has unit root	**KPSS test statistic** Null hypothesis: variable is stationary	**ADF type unit root test with structural break (Lanne et al. (2002)) test statistic** Null hypothesis: variable has unit root
$M2Y^{adj}$	Levels (constant)	-0.00 (0.95)	0.58**	-0.13
	Levels (constant and trend)	-6.78 (0.00)	0.14*	-0.29
	1st differences (constant)	-5.30 (0.00)	0.48**	-3.42**
Y	Levels (constant)	-2.00 (0.29)	0.44*	-1.26
	Levels (constant and trend)	-4.82 (0.00)	0.18**	-2.9*
	1st differences (constant)	-6.63 (0.00)	0.04	-3.24**
W	Levels (constant)	-0.99 (0.75)	0.42*	-2.2
	Levels (constant and trend)	-1.41 (0.85)	0.14*	-2.29*
	1st differences (constant)	-6.84 (0.00)	0.18	-3.14**
OC^{M1}	Levels (constant)	-7.06 (0.00)	0.41*	-3.56**
	Levels (constant and trend)	-5.49 (0.00)	0.16**	-0.2
	1st differences (constant)	-6.56 (0.00)	0.42*	-4.13**
OC^{M2}	Levels (constant)	-21.3 (0.00)	0.31	-1.65**
	Levels (constant and trend)	-20.0 (0.00)	0.16**	-0.81
	1st differences (constant)	-6.82 (0.00)	0.37	-4.75**
OC^{M2Y}	Levels (constant)	-2.83 (0.06)	0.31	-3.94**
	Levels (constant and trend)	-2.97 (0.15)	0.16**	-0.36
	1st differences (constant)	-8.15 (0.00)	0.06	-5.28**
UNC	Levels (constant)	-2.67 (0.09)	0.15	-3.09**
	Levels (constant and trend)	-2.56 (0.30)	0.15*	-1.29
	1st differences (constant)	-4.35 (0.00)	0.13	-2.58

** – rejection of the null at 5%-level, * – rejection of the null at 10%-level.

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Figure A2
Recursively estimated eigenvalues with 95% confidence bands (for fixed short run dynamics)

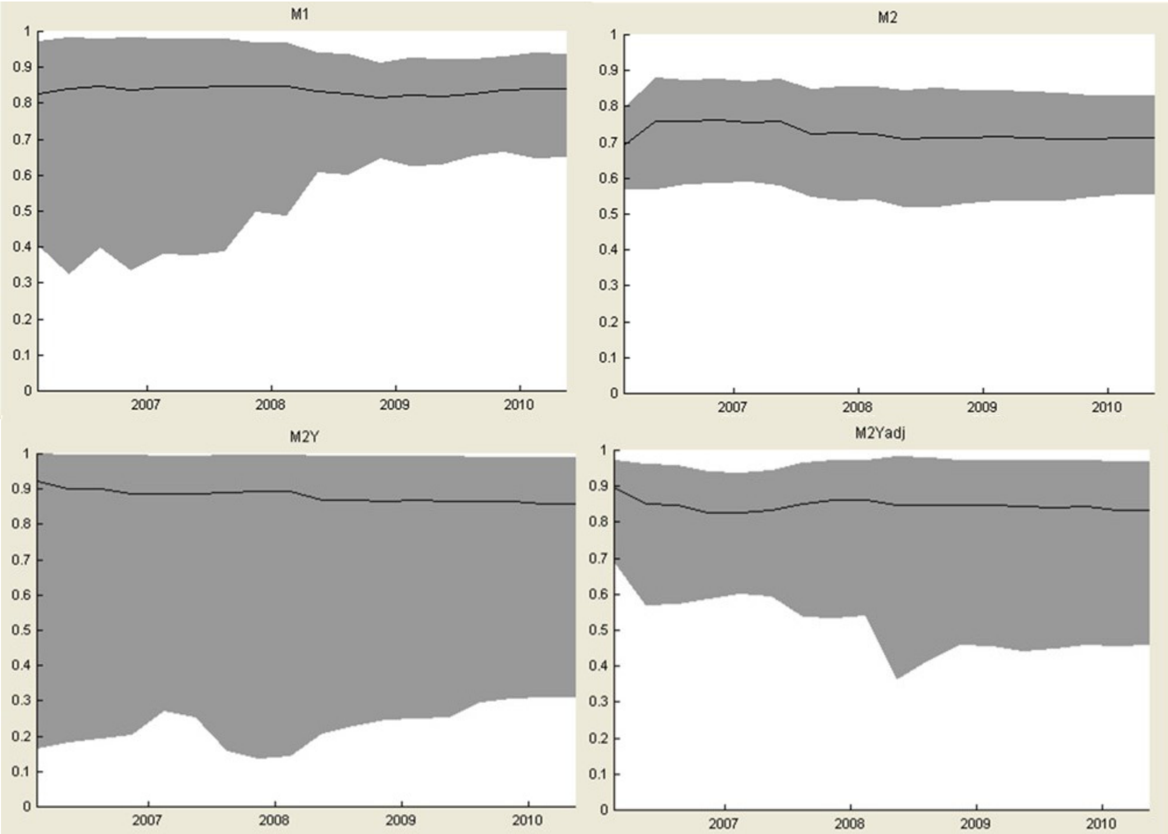

Figure A3
Recursively estimated statistic of Hansen and Johansen (1999) fluctuations test and 95% critical value (for fixed short run dynamics)

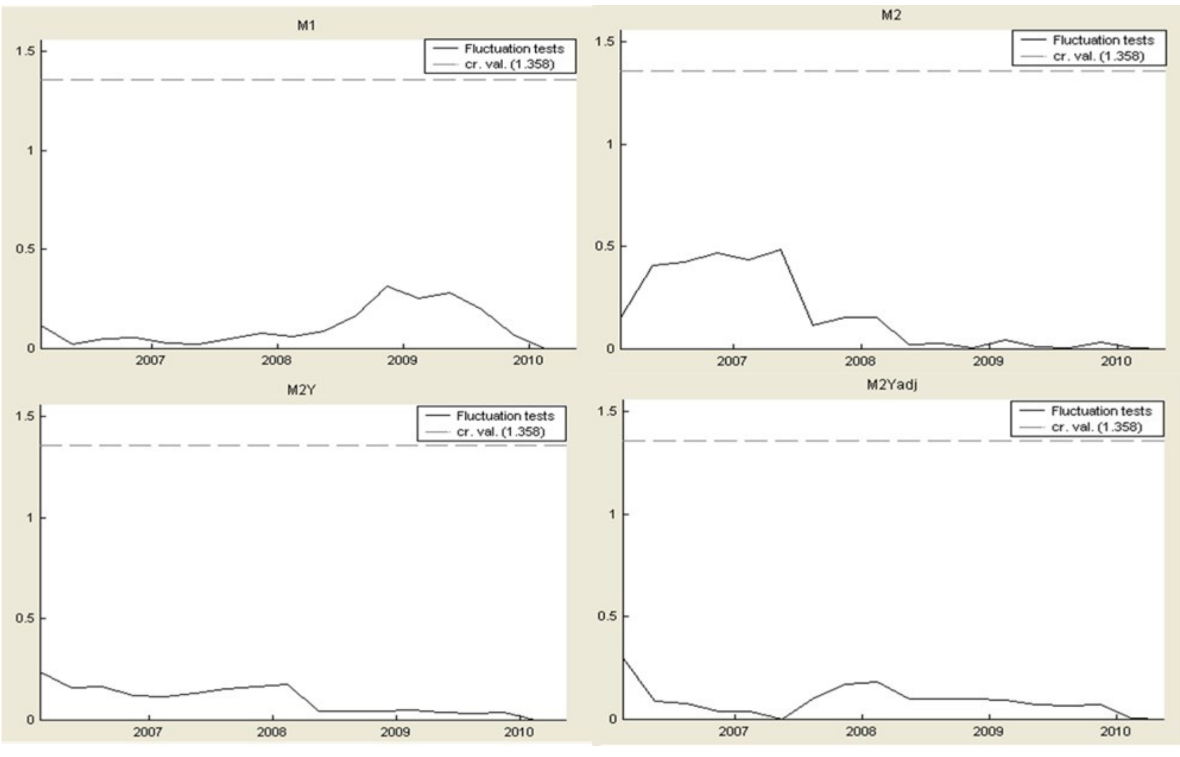

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Table A2
Test for stationarity of variables: F-statistics (p-value)

Variable	Model			
	M1	**M2**	**M2Y**	**M2Y^{adj}**
M	70.65 (0.00)	46.74 (0.00)	71.53 (0.00)	59.06 (0.00)
Y	71.25 (0.00)	46.55 (0.00)	70.06 (0.00)	60.13 (0.00)
W	67.71 (0.00)	44.33 (0.00)	69.03 (0.00)	53.54 (0.00)
OC	50.06 (0.00)	42.58 (0.00)	31.13 (0.00)	36.57 (0.00)
UNC	33.50 (0.00)	29.00 (0.00)	51.31 (0.00)	51.09 (0.00)

Figure A3
Recursive estimates of M1 money demand model's cointegration vector (±2 S.E.)

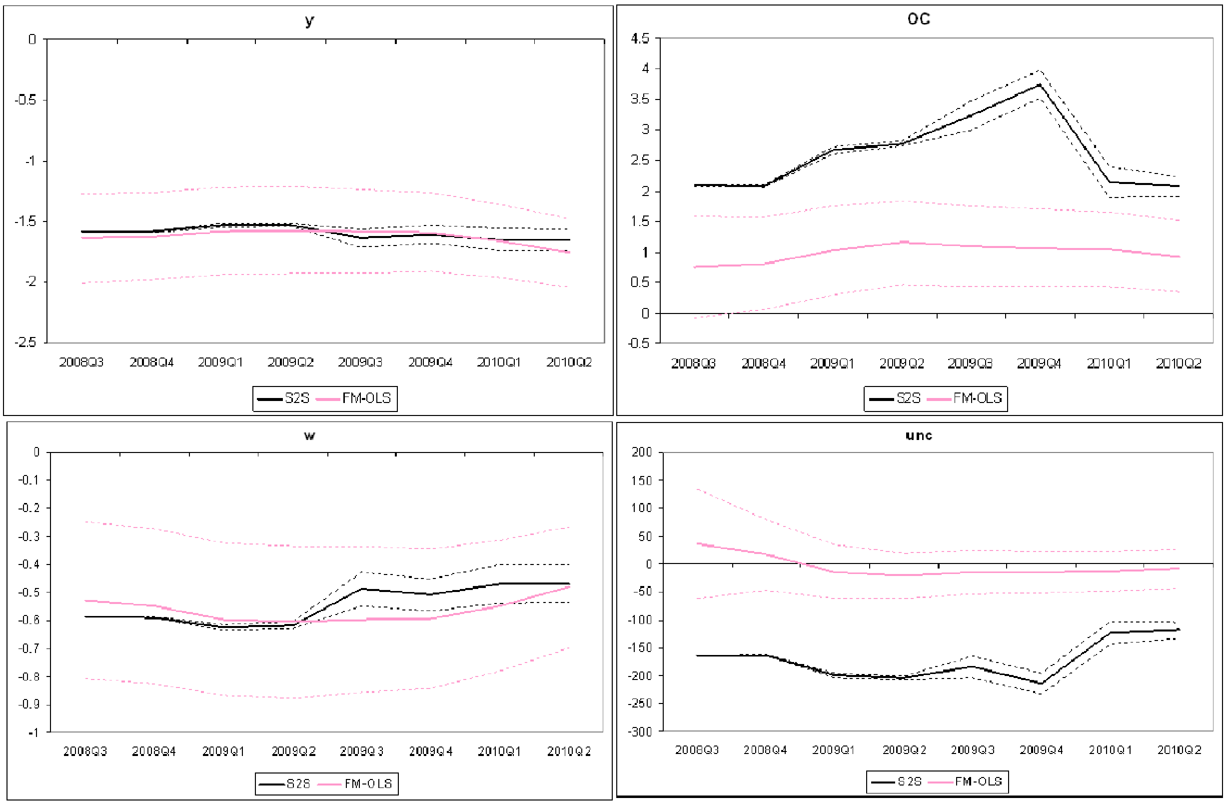

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Figure A4
Recursive estimates of M2 money demand model's cointegration vector (±2 S.E.)

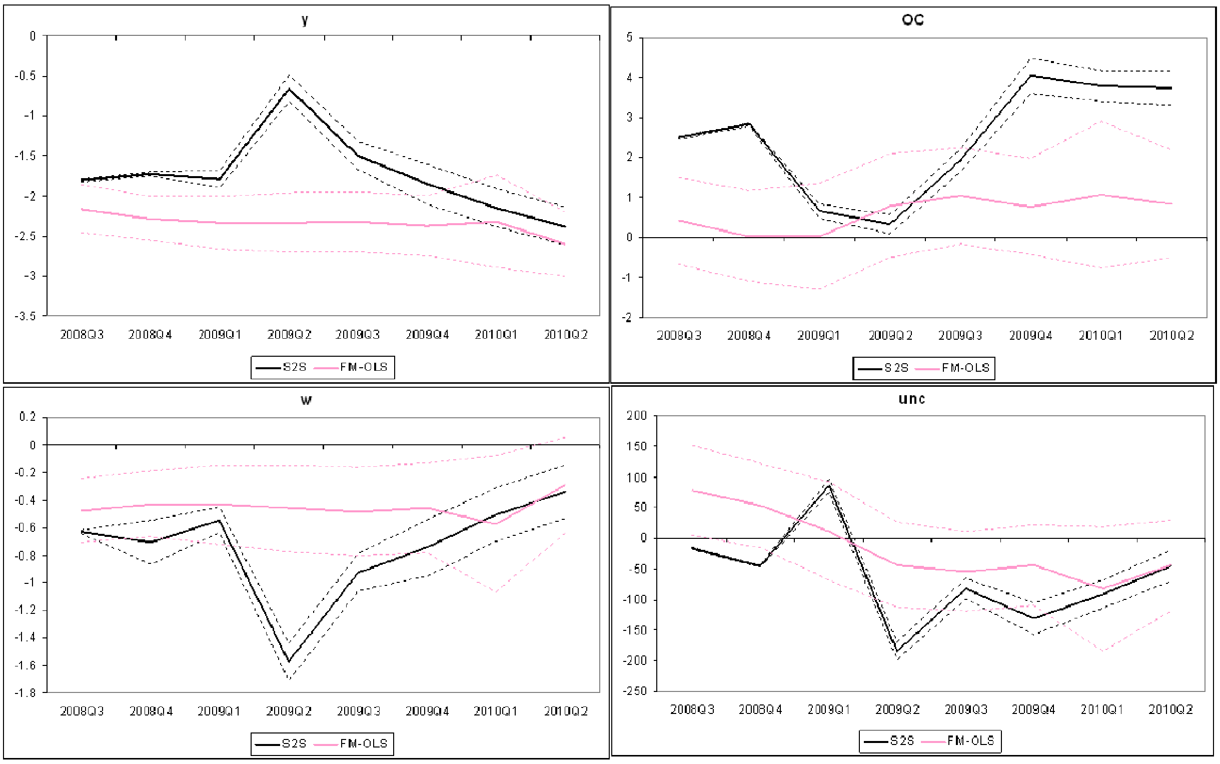

Figure A5
Recursive estimates of M2Y money demand model's cointegration vector (±2 S.E.)

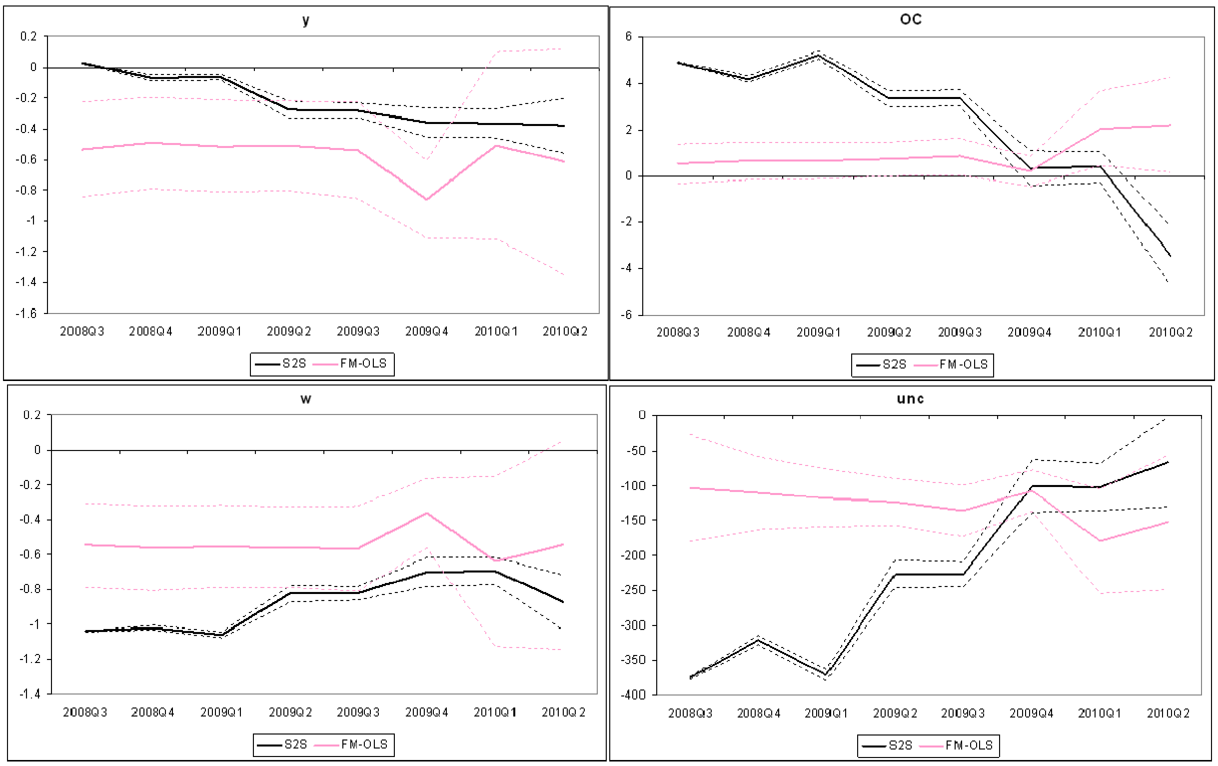

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Figure A6
Recursive estimates of M2Y^{adj} money demand model's cointegration vector (±2 S.E.)

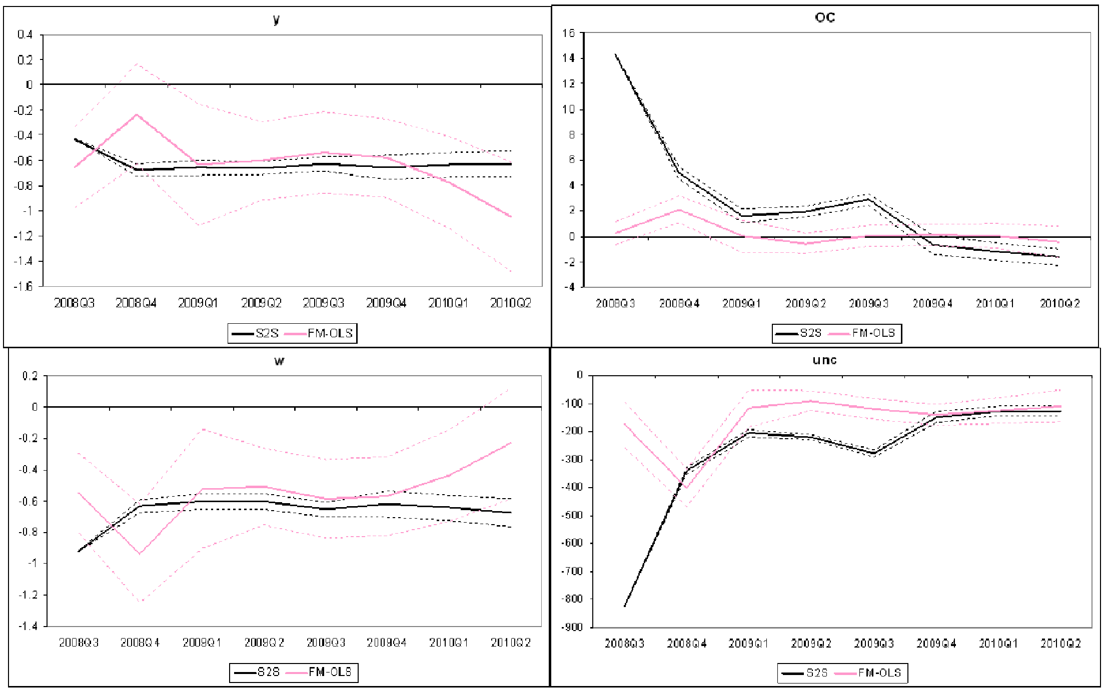
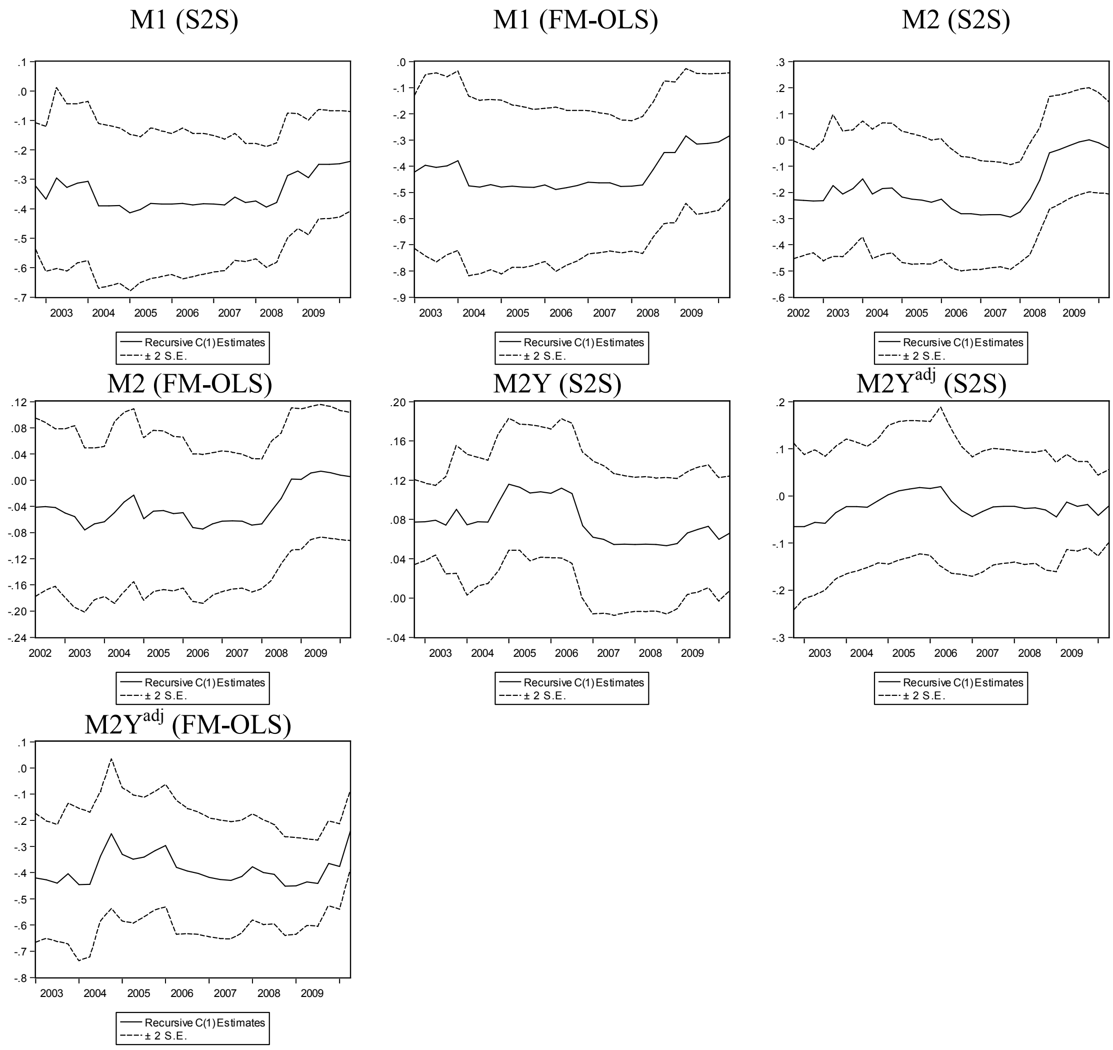


Figure A7
Recursive estimates of ECMs' loading coefficients



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ANNEX B

Figure B1
Variables used in equitation (5) (shaded area = period used for out-of sample forecasts)

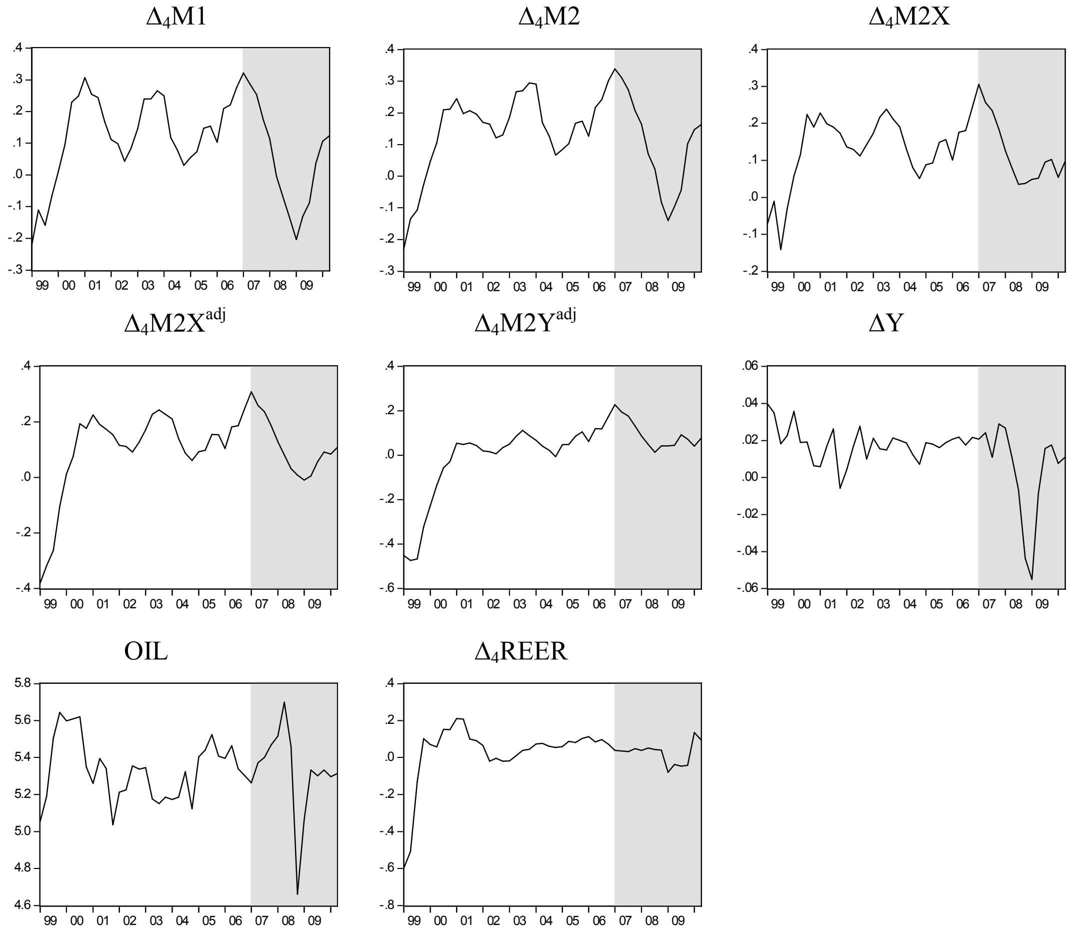

Table B1
Results of the unit root tests (variables in levels, bandwidth determined by automatic Andrews (1991) procedure)

Variable	**PP test statistic (p-value)** Null hypothesis: variable has unit root	**KPSS test statistic** Null hypothesis: variable is stationary
$\Delta_4 M1$	-1.99 (0.29)	0.31
$\Delta_4 M2$	-1.84 (0.36)	0.34
$\Delta_4 M2X$	-1.75 (0.4)	0.28
$\Delta_4 M2X^{adj}$	-3.55 (0.01)	0.35*
$\Delta_4 M2Y^{adj}$	-2.6 (0.1)	0.38*
ΔY	-4.55 (0.00)	0.23
OIL	-3.0 (0.05)	0.12
$\Delta_4 REER$	-5.48 (0.00)	0.24

* rejection of the null of stationarity at 10%-level.

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