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Essays in Small Open Macroeconometrics

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Contents

To Federica,
sister and nurse

Introduction

This thesis is composed by three papers connected to the study of the business cycle of small open economies under a double prospective: the consequences related to the transmission mechanism of global shocks and the employment of forecasting methods to date future economic developments at national level.

Specifically, the first chapter analyses the role of international shocks on a small non-oil commodity-exporting economy, using Australia as representative case. The Australian natural resource sector is independent from oil and this is a single feature among developed resource-rich countries. Identifying global shocks through a combination of sign and bounds restrictions, the results reproduce the main stylized features commonly documented in this type of economies, e.g. concerning the dynamics of commodity currency and terms of trade. Particularly interesting is the response of trade-related sectors after the rise in commodity price; indeed no Ducth disease is found even if the reaction of nontradable sectors testifies its function is essential for the transmission of beneficial effects of a global shocks on real domestic variables. Furthermore it is shown that for some Australian variables the source behind the movement of resource price is not relevant while for other indicators the origin of the shock must be distinguished.

The second chapter develops an early warning system for predicting incoming recessions for a group of small open economies. Combining models and data from previous

contributions on early warning systems and business cycle analysis in small open economies, the results turn out to highlight the presence of some commonalities among this type of economies in terms of signaling for recession outbreaks. Out-of-sample performance also confirms the ability of the model to signal for recession outbreaks. Despite this, the baseline model is tested with different business cycle identification and augmented with global factors outside national policymakers control, suggesting the dependence of the results to specific specification and in particular to recession duration. In the third paper the predictability of monthly inflation for three small open economies is assessed. Through the Bayesian estimation of a set of linear models, it is worth claiming that including a measure of global inflation to a standard AR(1) model for national inflation is not beneficial either in terms of point and density forecasts. This is then faced considering the dynamic of regional inflation together with the global one in the benchmark specification. This strategy suggests the importance of sub-global developments in driving domestic indicators; the results are consistent at point and density level and for two different prior specifications.

Chapter 1

Global shocks and small non-oil commodity-exporting economies: a Bayesian FAVAR model for Australia

Global shocks and small non-oil commodity-exporting economies: a Bayesian FAVAR model for Australia[*]

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Abstract

This paper estimates an open-economy Bayesian FAVAR model for Australia, which can be considered as a small commodity-exporting country. The sectorial composition of Australian economy represents farther a singular case among this type of countries for the small contribution of the oil sector to the overall GDP. Identifying global shocks through a combination sign and bounds restrictions, the results turn out to reproduce the main stylized facts about the effects of world developments on resource-rich economies. Particularly interesting is the role played by the tradable sector when commodity price boom takes place; indeed no beneficial effects are computed for real trade-related variables, while positive outcomes are depicted for nontradable sector and transmitted to the real state of the economy.

Keywords: Factor models; VARs; commodity price; business cycle; commodity exports.

JEL Classification: C32, Q02, E32.

1 Introduction

The constantly increasing international openness and the subsequent global integration of national economic activities are considerable reasons used to explain the impact of the fluctuations of world (macro) variables on domestic economic indicators. The recent 2008-2009 financial crisis highlighted the extreme easiness through which the shocks spread among and within the economic networks all over the globe. Obviously the latter example of worldwide connection is the most notable, research and media covered since its beginning; however, the economic system is always subject to structural shocks of several magnitudes which are needed to be studied and that are not widely known except for policymakers and researchers and which usually connote the behaviour of national economic variables even in the medium-long run. Since "not all oil price shocks are alike" (Kilian, p.1064, 2009), the same story can be told for other types of international shocks. Thus, it is useful to disentangle different structural shocks to the business cycle that arise in shaping dynamic fluctuations to portray the real effects of specific events at country level. Besides, especially in those countries whose economic activities are consistently internationally-oriented, which fall under the name of small open economies, foreign developments are key points which have to be dealt with to ward the internal economic processes. Particularly interesting and narrowly studied by the increasing literature of structural VARs is the influence of commodity market developments (different from oil-specific shocks) on this type of countries.

To this aim, this paper investigates the role played by commodity price, and the indirect channels through which it is stimulated, on a small commodity-exporting economy, namely Australia. Since previous contributions mainly focus on the effects of international shocks on oil-importing countries, fewer attentions have been devoted to exporters of primary commodities and specifically to exporters of commodities other than oil, which can be indicated as small non-oil commodity-exporting economies (hereafter SNOCEEs); Australia represents, indeed, a good example in this sense and, despite this feature, receives little consideration so far.

To consistently deal with the huge amount of data depicting global and domestic business cycle, the empirical framework is built on factor-augmented VAR (FAVAR) model introduced by Bernanke, Boivin and Eliasz (2005) and then further specified by Mumtaz and Surico (2009) and Boivin and Giannoni (2009). Model estimation is performed with a modern Bayesian procedure to control for possible parameters uncertainty and, straightforwardly, model instability. The standard identification issue connected to structural VARs framework is faced with the employment of sign and bounds restrictions imposed through to the efficient algorithm introduced by Rubio-Ramirez, Waggoner and Zha (2010).

As previously stated, Australian business cycle can be taken as a good example to draw new evidence about commodity-producing countries. In fact, the contribution of the oil sector to the overall GDP is tiny (0.17% in 2016) with respect to other natural resources industries (Figure 1) and the oil net exports are indeed negative; these represent a singular features among small commodity-exporting developed economies; then the business cycle statistics reported in Table 2[1] convey the reliance of the external accounts to commodity-related activities[2]; moreover the availability of data at global and national level makes the analysis implementable either for the number of time observations and for the quantity of indicators.

Understandably, the model is composed by a global and a country block, in line with previous contributions (among the others, Boivin and Giannoni 2009, Mumtaz and Surico, 2009, Charnavoki and Dolado, 2014). The choice to include worldwide variables instead of specific regional factors is to avoid mistaken outcomes in shocks identification. For example, even if China is the Australia main trading partner (33% of 2016 Australian total

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[1]The statistics about Australia reported in Table 2 are collected at yearly frequency and they cover the period from 1991 to 2016 due to data availability.

[2]As it is shown in Table 2 the relief of the commodity sector (agriculture, forestry, fishing and mining) is smaller with respect to the noncommodity tradable (manufacturing) sector and to the nontradable sector (construction) , in terms of overall GDP and share of total employment. The situation somewhat changes if the external sector and especially the trade balance (net exports) are observed; there is indeed a reverse effect on overall GDP, where the commodity sector represents a positive large share, while the manufacturing's net exports are a negative portion of the real activity, on average.

exports[3], according to the Observatory of Economic Complexity) and it may be treated as the main source of its foreign demand fluctuations, keeping other global indicators out of the model specification might omit to consider commodity market turbulence taking place in other areas, undermining the reliability of the results.

The main findings can be summarized as follows. First, the paper provides evidence about the existence of different channels other than specific shocks through which commodity prices are stimulated (in line with Kilian, 2009). Secondly, the standard response of a commodity currency after a surge in resource price (Chen and Rogoff, 2003 and Cashin, Céspedes, and Sahay, 2004) and its positive correlation with the terms of trade reaction is confirmed. Thirdly, since the evidence of the effect of a rise in commodity price on the external balances is almost completely restricted to oil-exporting countries so far[4], this finding is here highlighted for a SNOCEE; the results are in line with the ones obtained by Kilian, Rebucci, and Spatafora (2009) for importers of oil at aggregate level. Finally, still connected to the movements of commodity price, the source of the price increase is shown to matter to outline the real effects on Australian GDP and expenditure components. As for the last point, even if no Dutch disease effects is detected, the evidence is consistent with the ones by Spatafora and Warner (1999) and more recently by Bjørnland and Thousrud (2016) for resource-rich countries concerning the different speed of reaction of specific individual industries (in particular nontradable sector) in response to stimulation in traded sector. Relatively to this it is worth highlighting the fact that being oil price mostly stimulated by a global shock, nonpositive effects are produced by a resource price boom on commodity related Australian variables.

The paper is structured as follows. Section 2 provides a literature review. Section 3 describes the empirical model, the identification of the structural shocks, the estimation method and the data collected. Section 4 presents a battery of detailed results linked to the estimation of factors, identification of shocks and propagation mechanism on the SNOCEE. Section 5 concludes.

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[3]In general exports towards East Asian countries like China, Japan and South Korea represent more than 50% of Australia's total exports.

[4]As suggested by Charnavoki and Dolado (2014).

2   Literature review

The need for investigation of structural shocks had been initially met through the widely-employed VAR models (Burbidge and Harrison, 1985, Johnson and Schembri, 1990, Kuszzak and Murray, 1987 and Souki, 2008), which bring along the feature of the model scale strictness because of their inability to cover for a large number of variables. Furthermore, the applications have been related to the United States in most of the cases and linked to the identification of a variety of global business cycle or energy commodity price shocks, in particular oil (Kilian, 2009, Kilian and Murphy, 2012 and Lippi and Nobili, 2012). Kilian, Rebucci and Spatafora (2009) document the role of the oil in the trade balance in shaping fluctuations in response to sudden economic developments. More recent contributions exploiting non-linear VAR models are partly justified by the increasing number of works focused on economic policy uncertainty leading most of the attention to its role and its degree of affection on oil price (Van Robays, 2016). The willingness to better portray the dynamics of macro-shocks have been attested by the introduction of latent variables to frame the factor analysis. This has been straightforwardly used to advocate the empirical works about oil market and business cycle shocks, both under a comprehensive perspective. The different responses to several oil shocks (price, demand, supply) has highlighted the existence of distinguishable features between oil-exporting and oil-importing countries (Peersman and Van Robays, 2012) or emerging and developed countries (Aastveit, Bjørnland and Thousrud, 2015). Novel specification methods accompany the evolution of factor models (firstly submitted by Bernanke, Boivin and Eliasz, 2005): among the others, the specification which imposes structural dynamics of the foreign and domestic business cycle for small open economies (Mumtaz and Surico, 2009) to drive new applications in international macroeconomics fields. The introduction of ad hoc specified FAVAR model gave birth to the general set-up to be borrowed across several works and country case studies; e.g. Canada, a textbook example of small open economy, it is hugely employed to

analyze the effects of global fluctuations on domestic variables (among the others, Maier and Vasishtha, 2013 and Charnavoki and Dolado, 2014) and, for the same sake, also the United Kingdom (Mumtaz and Surico, 2009). Evidences from Canada convey its dependence to global demand and commodity price shocks, while weaker effects are drawn from global supply and monetary policy shocks. The choice of Canada, as well as of the United Kingdom, has been also driven by data availability and Australia represents a good opportunity to fill the gap among the small open economies under the lens, among the other topics, of commodity-related economic activities. This picking also lies on the fact that very little has been done to capture the effects of macroeconomic shocks on Australia hitherto, at least with dynamic factor models (DFMs); a latter-day work by Bjørnland and Thousrud (2016) account for different spillover effects among sectors via a Bayesian DFM when Dutch disease arises in Norway and Australia. Then other papers fall in the small-scale VARs context in which the influence of specific foreign areas and countries on the Australian economy is one-by-one estimated and pointed out (Dungey and Fry, 2003 and Dungey, Osborn and Raghavan, 2014) next to the comprehension of the contributions of external shocks in the forecast errors of the domestic variables; in particular they show the output growth occurring in the United States have a positive correlated spillover effect on the Australian business cycle with respect to output growth promenading from Europe, while inflation variables and interest rates do not show different reactions to other shocks calling for a wider consideration of the sources of shocks at global level; afterward, Knop and Vespignani (2014) look for the responsiveness of Australian GDP sectors to commodity price shocks: they significantly estimate increasing profits in the construction, mining and manufacturing industry subsequent to a sudden development in the commodity market. VARs are sometimes accompanied by New Keynesian small open economy models (Nimark, 2009); Bhattacharyya and Williamson (2011) describe the reactions of Australia to terms of trade fluctuations due to exports prices volatility shocks; they also highlight the importance of stable revenues coming from different sectors of the economy which allow for controlled responses to exports shocks. A fresh work by Aastveit, Bjørnland and Thorsrud (2016) includes regional factors in the DFM, besides

the global factors, to control for sub-global developments in the country's main trading
partners and their economic geographical area.

3   Empirical framework

3.1   FAVAR model

As already anticipated in the introduction, the empirical model is built upon two-blocks
FAVAR model (Boivin and Giannoni, 2009 and Mumtaz and Surico, 2009): the first
outlines the global business cycle while the second refers to the SNOCEE. A number of
latent factors are extracted from a panel of series related to the two blocks and they aim
at empirically covering the main developments which occur along the sample period. In
particular, the two blocks of latent variables are $(\mathbf{F}_t^{W'}, \mathbf{F}_t^{D'})$ where the superscripts W
and D indicate world and domestic economy, respectively. On one hand, the vector of the
global economy, $\mathbf{F}_t^{W'}$, comprises four world factors and one world series, namely a world
economic activity factor, $F_{Y,t}^{W'}$, which captures the occurrences of the global economic ac-
tivity; a commodity price factor, $F_{C,t}^{W'}$, describing the global commodity market activity;
a global inflation factor, $F_{\Pi,t}^{W'}$, which depicts the dynamic of world price level; a global
liquidity factor, $F_{L,t}^{W'}$, denoting information about the money circulation on global scale;
the United States federal interest rate, $R_t^{W'}$, is the only directly observable series entering
the global block and it is used to control for the level of the global monetary policy. On
the other hand, the vector of the SNOCEE, $\mathbf{F}_t^{D'}$, is built on a number of factors extracted
from national series and employed to get a reliable picture of the Australian business
cycle. Specifically, three domestic factors are extracted; this quantity seems plausible and
in line with previous contributions (e.g. Aastveit, Bjørnland and Thorsrud, 2016) and
they do not need any economic interpretation or distinguishing identification.

The model is composed by two equations: an observation equation which describes the
relation between the unobserved factors and the different sets of variables and a transition

equation, which connotes the dynamics of the latent factors. Respectively

$$
\begin{pmatrix} \mathbf{X}_{Y,t}^W \\ \mathbf{X}_{C,t}^W \\ \mathbf{X}_{\Pi,t}^W \\ \mathbf{X}_{L,t}^W \\ \mathbf{R}_t^W \\ \mathbf{X}_t^D \end{pmatrix} = \begin{bmatrix} \Lambda_Y^W & 0 & 0 & 0 & 0 & 0 \\ 0 & \Lambda_C^W & 0 & 0 & 0 & 0 \\ 0 & 0 & \Lambda_\Pi^W & 0 & 0 & 0 \\ 0 & 0 & 0 & \Lambda_L^W & 0 & 0 \\ 0 & 0 & 0 & 0 & \Lambda_R^W & 0 \\ \Lambda_Y^D & \Lambda_C^D & \Lambda_\Pi^D & \Lambda_L^D & \Lambda_R^D & \Lambda^D \end{bmatrix} \begin{pmatrix} F_{Y,t}^W \\ F_{C,t}^W \\ F_{\Pi,t}^W \\ F_{L,t}^W \\ R_t^W \\ \mathbf{F}_t^D \end{pmatrix} + \begin{pmatrix} \mathbf{e}_{Y,t}^W \\ \mathbf{e}_{C,t}^W \\ \mathbf{e}_{\Pi,t}^W \\ \mathbf{e}_{L,t}^W \\ \mathbf{e}_{R,t}^W \\ \mathbf{e}_t^D \end{pmatrix} \tag{1}
$$

and

$$
\begin{pmatrix} \mathbf{F}_t^W \\ \mathbf{R}_t^W \\ \mathbf{F}_t^D \end{pmatrix} = \begin{bmatrix} \beta_F^W(L) & 0 & 0 \\ 0 & \beta_R^W(L) & 0 \\ \beta_{W,F}^D(L) & \beta_{W,R}^D(L) & \beta_F^D(L) \end{bmatrix} \begin{pmatrix} \mathbf{F}_{t-1}^W \\ \mathbf{R}_{t-1}^W \\ \mathbf{F}_{t-1}^D \end{pmatrix} + u_t \tag{2}
$$

In equation (1) X_t^W and X_t^D refers to global and domestic series; F_t^W and F_t^D indicates the latent factors extracted from world and Australian data which are related to the observable series through the loading Λ^W and Λ^D; e_t^W and e_t^D are measurement errors such that $E(e_t^W) = E(e_t^D) = 0$ and $cov(F_t, e_t) = 0$.

The restricted structural VAR in equation (2) consists of latent factors, its lags up to finite order p loaded by respective lag polynomials $\beta_{i,j}^z(L)$ and reduced-form residuals u_t which are assumed to be $u_t \sim N(0, \Omega)$ with $u_t = B_0 \epsilon_t$, where $\Omega = B_0 B_0'$ and ϵ_t is a normally distributed structural shock with zero mean and variance I. A couple of things need to be specified. First it is useful to recall that the global interest factor corresponds to the global interest rate series itself; second it is worth noting that the global factors are included in the last row of equation (1) to explicitly be considered as, together with

the domestic factors, drivers of the SNOCEE business cycle.

In light of what is specified in this first block and of the data chosen, described in the next subsection, it may be appropriate to qualify that the consideration of these international factors aims combine the specification employed by Mumtaz and Surico (2009), for what concerns the introduction of a liquidity factor next to the global interest rate, and Charnavoki and Dolado (2014), with respect to the inclusion of a commodity price factor. However some additional motivations could be necessary. Relatively to the first two factors now named, on the one hand their dynamic might be thought as depicting equivalently a monetary shock. On the other hand they can be seen under two different aspects of monetary policy: a liquidity shock can be interpreted as the trigger of an open market operation with the aim to increase the money supply in a country or a specific economic area (e.g. the European Union) through the purchase of government bonds by the Central Bank, for example; while a shock in the interest rate factor might be connoted as instrument to accommodate such expansionary monetary policy to make it effective. The commodity factor does not need any further explanation concerning the decision of its inclusion in the model: it appears useful to the purpose of the paper to consider a specific dimension for the commodity market.

3.2   Data

In this section details about the data employed are given. The data are collected at quarterly frequency and they span the period comprised between 1981:I and 2016:II; a total of 93 series is gathered from different databases (International Monetary Fund, World Bank, OECD, FRED, Australian Bureau of Statistics and Reserve Bank of Australia) independently accessible or through the use of Datastream. The data depicting the global block refer to the international economy via world and country aggregates (OECD, European Union and G7) next to single economies large enough to impact the global business cycle, i.e. the United States. As shown in the empirical model section, the series collected for

the foreign block allow to extract four factors; the real activity factor is extracted from data about GDP, industrial production index, exports and imports; the commodity price is obtained from a group of five commodity aggregates; the global inflation is measured by consumer price indices and GDP deflators; several monetary aggregates are summarized by the global liquidity factor. The interest rate included in the model to control for global monetary policy is simply the United States federal funds rate (due to the short availability of data on world interest rates, especially in terms of time observations). Data on Australia are not collected according to the 'the more the best' logic; as suggested by Boivin and Ng (2006) and Caggiano, Kapetanios and Labhard (2011) the most representative (to my knowledge) series featuring and capturing the Australian state of economy are chosen; they range from real activity indicators to consumer and producer prices and interest rates, presumably allowing for a reliable approximation of the domestic business cycle. Non-stationary variables are differenced and all the variables are demeaned and standardized before proceeding with the estimation.[5]

3.3   Estimation strategy

The estimation procedure is conducted through a two-step principal component analysis, where in the first step the main principal component is extracted from the global and domestic series and then the factors extracted are employed in the restricted VAR. The two-step principal component analysis is borrowed from previous works by, among the others, Boivin and Giannoni (2009) and Mumtaz and Surico (2009). In particular, in the first step an iterative procedure is performed: starting from an initial estimate of the domestic principal component F_t^D, extracted from the domestic observed series, denoted by $F_t^{D,0}$, the iteration procedure advances as follows:

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[5]The variables are taken in log difference $[\log(X_t) - \log(X_{t-1})]$; only the trade balance (as % of GDP), the current account balance (as % of GDP) for the Australian block and the interest rates at global and domestic level are not differenced or taken in log.

1) Regress the domestic variables X_t^D on $F_t^{D,0}$ and estimates of the global factors to obtain the foreign and domestic loading matrices;

2) Compute $X_t^{D,0}$ as the difference between each domestic series and the loading matrices times the respective foreign factor, as estimated in the step above;

3) Estimate $F_t^{D,1}$ as the first principal component of $X_t^{D,0}$;

4) If necessary, go back to the first step to reach convergence in $F_t^{D,j}$.

This approach allows verifying that the international factors are truly common components since they should be captured by the principal component of the domestic series. This is connected to the model specification where the international factors are imposed to be included in the principal component for the Australian block of the model. The second step of the procedure concerns the estimation of the restricted VAR with 8 variables (5 global factors and 3 domestic factors). I proceed employing a Bayesian technique in order to deal with a high number of free parameters and observations. Moreover, Bayesian inference can efficiently manage the issue connected to uncertainty which can drive the parameters in the model. In particular, as in Charnavoki and Dolado (2014), I follow Koop, Poirier and Tobias (2007) for a likelihood estimation by multi-move Gibbs sampling; the parameters of the model are estimated by alternatively sampling from conditional posterior distributions (additional details are provided in the Appendix). The number of lags $(p = 1)$ necessary is controlled to allow to adequately capture the dynamics of the factors; a diagnostic test to verify the absence of residual serial correlation is performed.[6]

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[6]AIC criterion is performed to choose the lag for the restricted VAR; 1 lag is enough to capture the dynamics of the factors. This is also confirmed by the absence of serial correlation in the residuals (Ljung-Box test).

3.4 Identification of shocks

The identification of shocks always represents an issue to convincingly discuss. The procedure here adopted relies on the employment of the rotation procedure which allows to impose sign restrictions on the estimated impulse-response functions (Rubio-Ramirez, Waggoner and Zha, 2010); the procedure starts drawing B_0 through the Choleski decomposition, where the parameter matrix B_0 is taken to be lower triangular, and then advances up to creation of candidate impulse-response functions which are discarded if they do not satisfy the sign imposed; the signs are displayed in Table 1 and they are supposed to be respected along the first quarter. The problem associated to this identification procedure is that it does not produce point estimates of the impulse-response functions but instead it generates results derived from several different structural models. A further restriction is thus imposed; the functions which do not fail the test of signs are then evaluated according to the limitations to which the elements of the B_0 are constrained; in particular, following Kilian and Murphy (2012), a small short-run elasticity (between -10% and 10%) of the global real activity to a commodity price shock is assumed; hence the disrespectful results to this additional regard are rejected.

4 Results

The results are here reported. The main findings are organized to provide first a picture of the state of global economy with particular focus on the commodity market. Secondly the outcomes about the transmission mechanism of international shocks on a SNOCEE are shown.

4.1 Estimated principal components

The first step of the empirical analysis is connected to the observation of the principal components. The aim is to check for the power of the identified factors to depict the

Table 1: Sign restrictions imposed on the impulse-response functions

	Demand	Commodity	Supply	Liquidity	Monetary
Real activity	+	-	-	+	+
Commodity price	+	+	-	+	-
Inflation	+	+	+	+	+
Liquidity	-	+	NA	+	+
Interest rate	+	NA	NA	NA	-

Notes: The table shows the response of each global variable (in rows) to global shocks (in columns). NA indicates that no restrictions are imposed on the response of the variable to the corresponding specific shock.

state of the global economy along the sample period. The factors are plotted in Figure 2. The factors are shown to match the most important economic events as spotted by Kose, Ayhan, Otrok, and Whiteman (2003) and Mumtaz and Surico (2009). In particular they pick the main global occurrences strictly associated to the real economic activities, namely: the recession in the first half of the 1980s, the early 1990s recession linked to the Dow Jones Industrial Average collapse, the Exchange Rate Mechanism crisis in 1993, the East Asian financial crisis in the second half of the 1990s, the downturns triggered by the Dot-Com bubble and then by 9/11 terrorism attack in the early 2000s and the Great Recession triggered by the financial crisis started in 2008. In line with Boivin and Giannoni (2009) the factors also possess their own features. For example, the Great Recession is the main downturn for the global economic activity factor; the global inflation factor shows a strictly decreasing path since the 1980s; the global liquidity factor appears to be strongly volatile in the first half of 2000s (2002-2003).

4.2   Contribution to commodity price factor volatility

The opinion about the existence of a single source behind the trigger of shocks in the international panorama might be seriously misleading. The contribution to the development of the volatility of a single world dimension can indeed be due to a variety of different factors; this prediction may be extended to the case at hand. Commodity market is as a matter of fact a global player in shaping dynamic fluctuations from either a national or a supranational point of view. Because of this, the commodity price volatility ought to be investigated to control for factors other than the commodity price to which fluctuations in the commodity market can be attributed, as shown by Kilian (2009). Figure 3 is plotted is order to check for this contingency. As it is easily noting, the sources generating commodity price movements are different from being unique. While the global monetary policy and the global liquidity shocks seem to play no prominent roles, global real demand (consistently with previous studies, e.g. Juvenal and Petrella, 2015) and global supply, together with commodity price shocks appear to be the main contributors to commodity price volatility development. As for Figure 2, the shaded areas represent the principal world economic occurrences taken place along the sample period; the choice to plot them is to highligth the fact that the main contributors to the commodity price factor volatility hold either during economic downturns and in "normal" times. This finding helps to move toward deepening the analysis whether and to what extent commodity price shocks and its source empirically matter inside a SNOCEE.

4.3   International factors response to global shocks

The evidence of the previous section force to enlarge the consideration of commodity price movements to the ones generate by global demand shocks and global supply shocks together with specific commodity price shocks. Before investigating their effects on the Australian economy, their impact on the global business cycle is preliminary detected. Figure 4 reports the impulse-response functions of the global factors reacting to the three

shocks considered.[7] It is immediate to see that a commodity price shock is the less effective with respect to the other two; its effects are positive on commodity price and on global inflation, while in the other three factors (global real activity, liquidity and interest rate) the responses are not statistically significant. Increase in global demand is positively correlated with reactions of the whole international economy, except for global liquidity which looks unaffected. The main reactions of the global business cycle are undoubtedly driven by the unexpected surge in global inflation; global supply shock, by the way, has a generalized negative effect on the world factors, namely global real activity, commodity price and interest rate, while no influence is estimated for global liquidity. For completeness the effects of the three main global shocks on the single observed series of commodities are also shown; the shocks are standardized such that they all result in the increase of one standard deviation of the commodity price factor on impact; Figure 5 clearly outlines the hump-shaped response of the oil price, as in Kilian, Rebucci and Spatafora (2009), to a world shock. The other commodities (especially mineral and industrial commodities) appear significantly impacted too, even if tenuously compared to oil.

4.4 The effects of global shocks on Australian economy

The discussion about the transmission of the shocks is now dealt with respect to its effects on Australian economy. As above, the shocks are standardized to one standard deviation increase of the commodity price factor on impact.[8] Following Charnavoki and Dolado (2014), the results are shown according to the relevance of the source behind the increase of the commodity price for several domestic indicators. In particular the impulse-response functions are grouped (i) when the source of the commodity price surge does not matter and (ii) when the increase of commodity price has different effects depending on the origin

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[7]The impulse-response functions are reported with its median value together with 68% confidence bands.

[8]Since also the global supply shock is standardized to one standard deviation increase of the commodity price factor on impact, the effects at global level are expected to be reversed with respect to the ones indicated in the previous subsection.

of the fluctuation.

4.4.1 Domestic response when the origin of commodity price surge is not relevant

No particular significance for the origin of the shock is computed for the terms of trade, the exchange rate and the trade balance variables (Figure 6). When the commodity price is high after a commodity price or a global demand or a global supply shock, the terms of trade improves and the reaction is quite persistent. The rise in the terms of trade is positively correlated (with regard to sign and persistence) with the appreciation of the exchange rate of commodity currency, drawing a typical result inside a commodity-exporting economy (Chen and Rogoff, 2003 and Cashin, Céspedes, and Sahay, 2004): the commodity currencies are indeed strongly correlated with the price of exported commodities, to which the rise in total export prices is mainly due. The result about the trade and the current account balance for an oil-importing country is new in some sense, given that the evidence for this effect is drawn for oil-exporting economies so far; the outcomes about the external balances replicate however the findings of Kilian, Rebucci and Spatafora (2009), which work aggregately with oil-importing countries: a commodity price increase leads to a temporary deterioration of the trade balance and of the current account balance; here it is clear that global demand driving up domestic demand is not beneficial for trade-related real variables (e.g. trade balance); the increase of Australian prices combined with the surge of internal demand[9] produces negative effects on real external balances. An overall observation of the impulse-response functions in Figure 6 sustains the greater effectiveness of a global supply shock with respect to the others either at domestic level.

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[9]Increase in domestic demand driven by stimulating global environment can generate surge in domestic demand for oil, which Australia imports on net basis.

4.4.2 Domestic response when the origins of commodity price surge need to be distinguished

The story is different for SNOCEE GDP and expenditure components: here is the case to strictly distinguish the source of the surge in commodity price since the shock behind generates different dynamics in this subset of national variables. Spatafora and Warner (1999) verify the presence of a positive correlation between the increase of the terms of trade and the reaction of consumption and investment in a resource rich country after a rise in the price of commodities. In Figure 7 the impulse-response functions describing the effects on spending aggregates are reported. At first glance the strenght exerted by a global supply supply shock (Panel C) is notably greater than the one exerted by a global demand (Panel B) and, especially, by a commodity price shock (Panel A). The latter, precisely, has almost no effects on the variables here considered; it only affects real personal consumption but significantly just on impact, while the worsen of real GDP is totally not significant. The surge in global demand symbolizes the economic stimulus coming from the international environment which has unsurprising positive effects, particularly, on imports; increase in world demand translates also domestically into (slightly) positive results for real GDP, real personal consumption and investment. Major change in spending aggregates is due, as expected, to global supply shocks. Furthermore it is significantly beneficial to all the indicators considered with an evident persistence on the trade variables (imports) and investments. It is useful to underline that the positive outcomes of real variables (e.g. real GDP) are not driven by trade indicators (exports in this case); indeed, in line with the previous set of results, the deterioration of the trade balance after a commodity price boom translates into a contraction of the tradable sector; hence it is plausible to think that global demand and supply shock are beneficial to GDP and investment through other components of the Australian economy (e.g. nontradable and noncommodity sectors).

The last slot of results has been intended to investigate the presence of possible Dutch disease effects inside the SNOCEE (Figure 8). Admittedly this effect is not detected but,

as shown by Spatafora and Warner (1999) and Bjørnland and Thousrud (2016), a resource price boom which leads to the appreciation of the exchange rate and the improvement of the terms of trade translates into faster increase of nontradable with respect to tradable output. It is worth saying that the outcomes at disaggregated level reflect the findings provided for the real overall GDP. Indeed no effect is detected on GDP sectors after a commodity price shock (Panel A); this is consistent with the evidence about the effects that a generalized surge in the price of commodities have on single commodities aggregates: the output in the Australian mining sector does not correlate with the increase in the price of oil triggered by a commodity price shock because of the small contribution of fuel to the overall GDP. This is even more true to the other commodity sector, namely the agriculture, forestry and fishing industry. The situation changes when the commodity price rises after a global demand and supply shock (Panel B and Panel C, respectively), at least for noncommodity and nontradable sectors. Manufacturing industry (noncommodity) is indeed stimulated mostly by the foreign and partially by the internal demand. However the domestic economy, as previously seen, is much more boosted when a world supply shock takes place: the revenues in either manufacturing and service sector (nontradable) are reaped as benefits promenading from the economic expansion and the persistence of the reaction configures long term gains.

5 Conclusions

This paper estimates a Bayesian FAVAR model for a small non-oil commodity-exporting economy, using Australia as representative case. The motivation behind the choice of analyzing the Australian economy is due to the composition of its real activities, in particular the structure of the natural resource sector. The main focus is related to the comprehension of the sources responsible for commodity price movements at world level and their effects on this type of economies. Employing a sign restriction identification scheme, the results support the predictable evidence of different sources behind resource price surge

other the ones strictly connected to commodity market. In particular global demand shocks and global supply shocks are assessed to strongly impact the price of commodities when structural developments take place in international business cycle. The relevance of the source of shocks is therefore valuated at domestic level, providing outcomes in line with previous contributions about resource-rich countries and now applicable also to Australia. Overall it is found that (i) a primary commodities price increase is not beneficial for real trade-related domestic variables and (ii) a global supply shock leading to a rise in price of commodities represents the main driver of domestic response to an international occurrence. Moreover the results testify that no significance has the source of commodity price surge when analyzing its effects on external balances and exchange rate; while the origin of commodity price movements becomes relevant when the reactions of GDP and expenditure aggregates and individual industries are detected. A key point to be stressed is the one indeed connected to the response of specific economic sectors; in fact, even if no Dutch disease effect is found, the findings about sectorial output confirm prior evidence on commodity-exporting economies related particularly to the positive correlation between nontradable sector outcomes and commodity price surge driven by terms of trade increase.
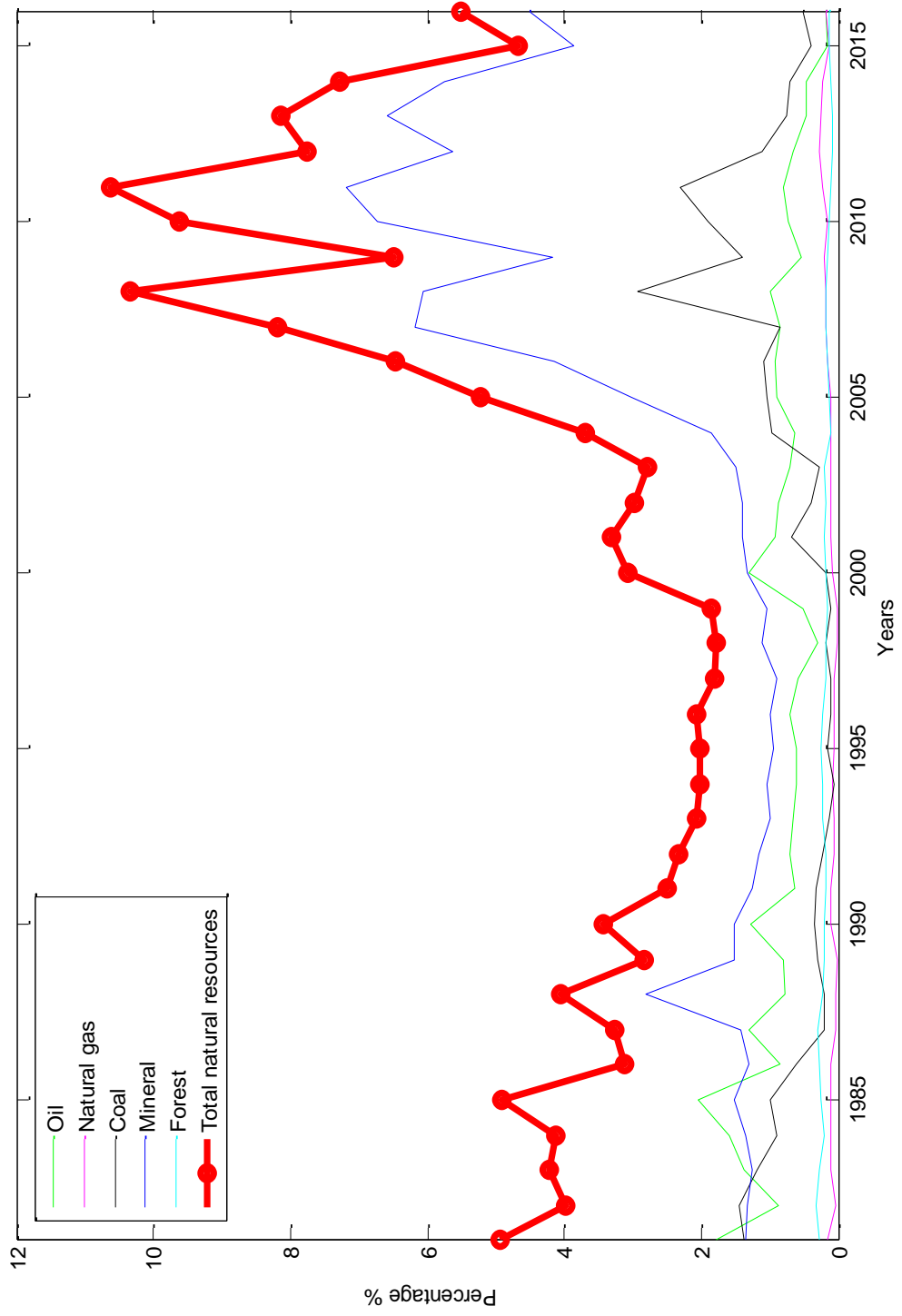
Figure 1: Composition of Australian natural resource sector (% of GDP).

Source : World Bank.

Table 2: Sectorial composition of Australian economy (1991-2016, yearly).

	Value added (% of GDP)	Employment (% of total employment)	Exports (% of total exports)	Net exports (% of GDP)
Agriculture, forestry and fishing	2.80	3.99	9.63	0.29
Mining	4.91	-	34.37	1.17
Manufacturing	6.79	8.07	-	-
Construction	10.25	11.68	52.02	-0.92

Notes: The table reports the average annual values from 1991 to 2016 (Source: Australian Bureau of Statistics).
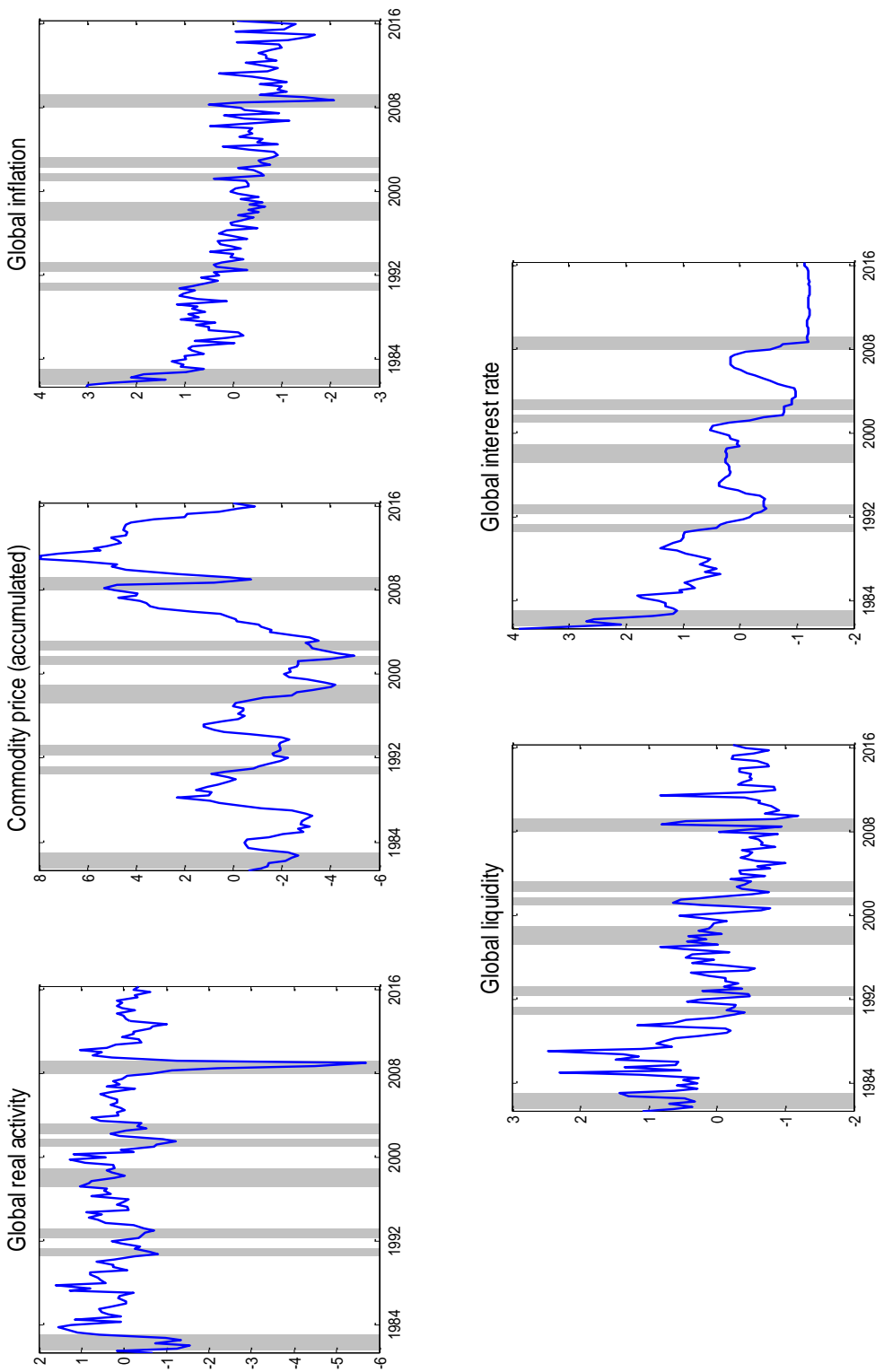
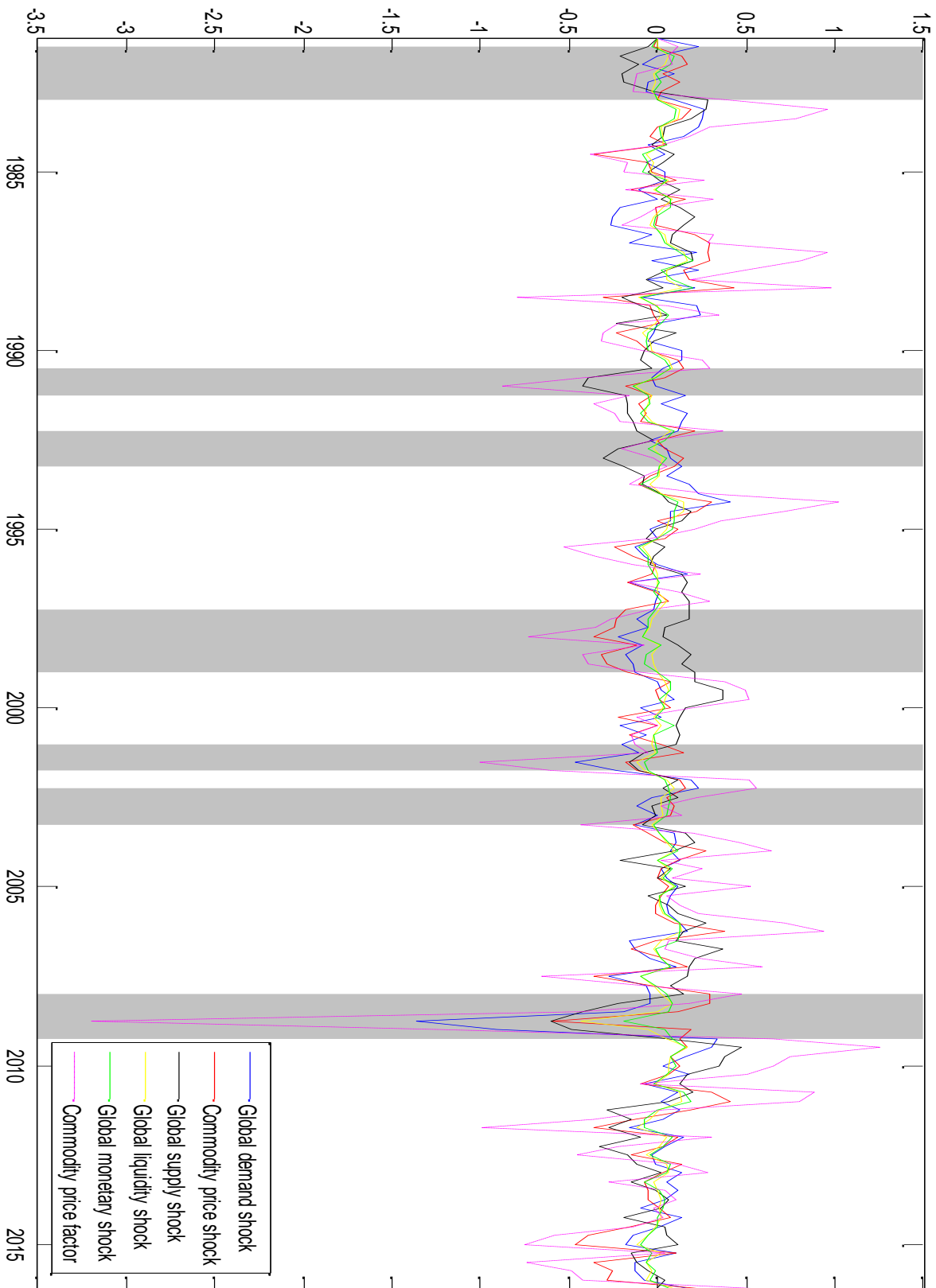Figure 2: Global factors (shaded areas indicate global recessions).

Figure 3: Contribution to commodity price factor volatility (shaded areas indicate global recessions).
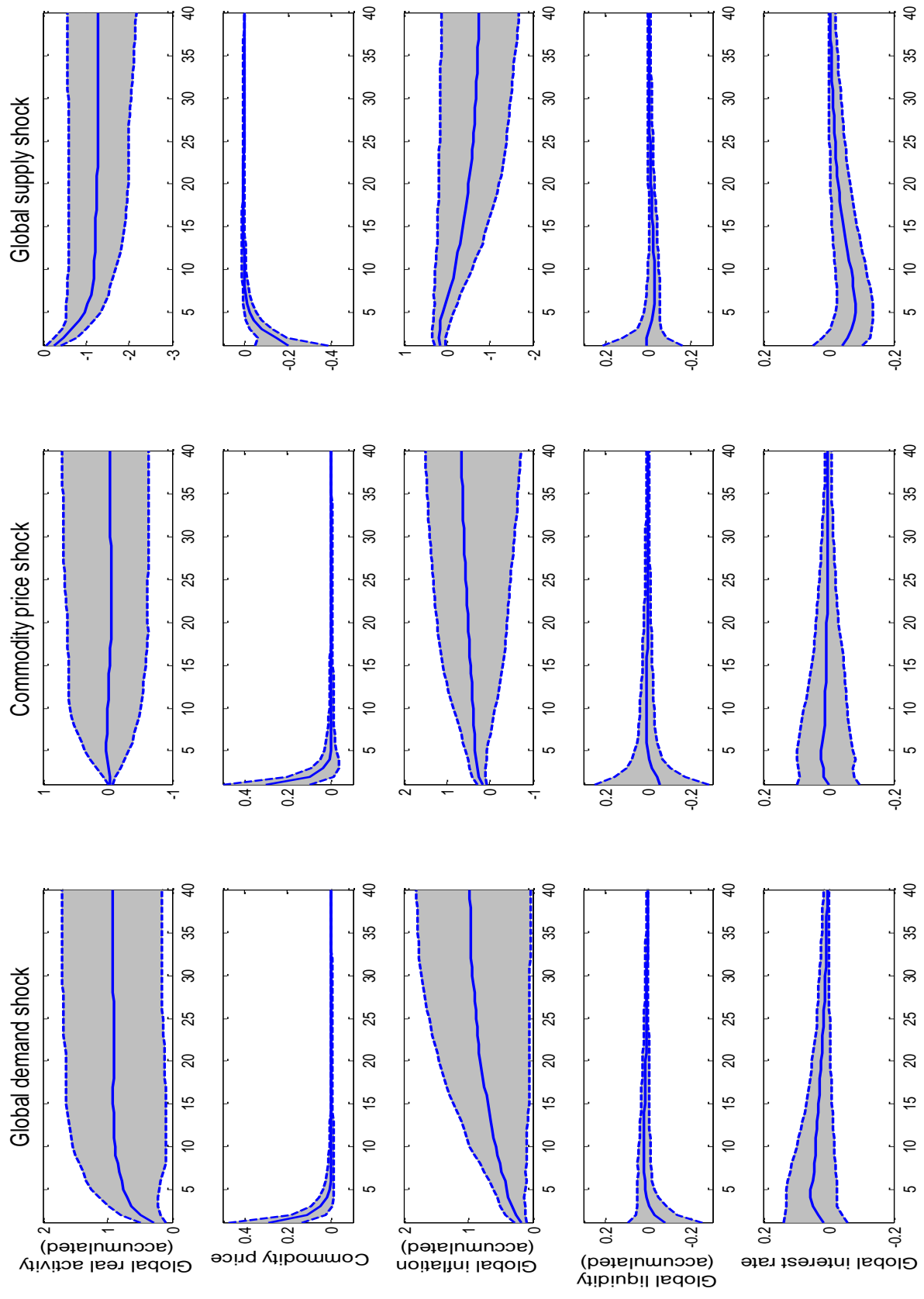
Figure 4: Impulse-response functions of global variables to the main international shocks. 68% confidence bands.
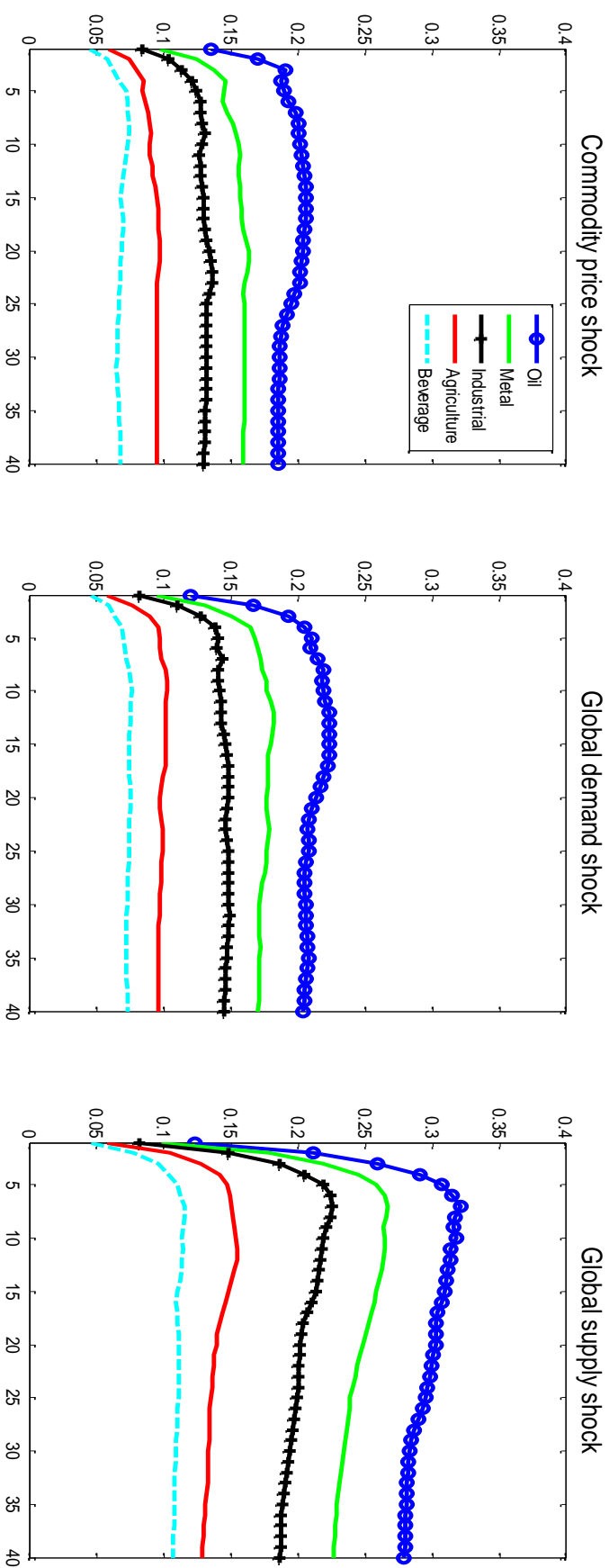
Figure 5: Impulse-response functions of commodity price factor components to global shocks. Median response.
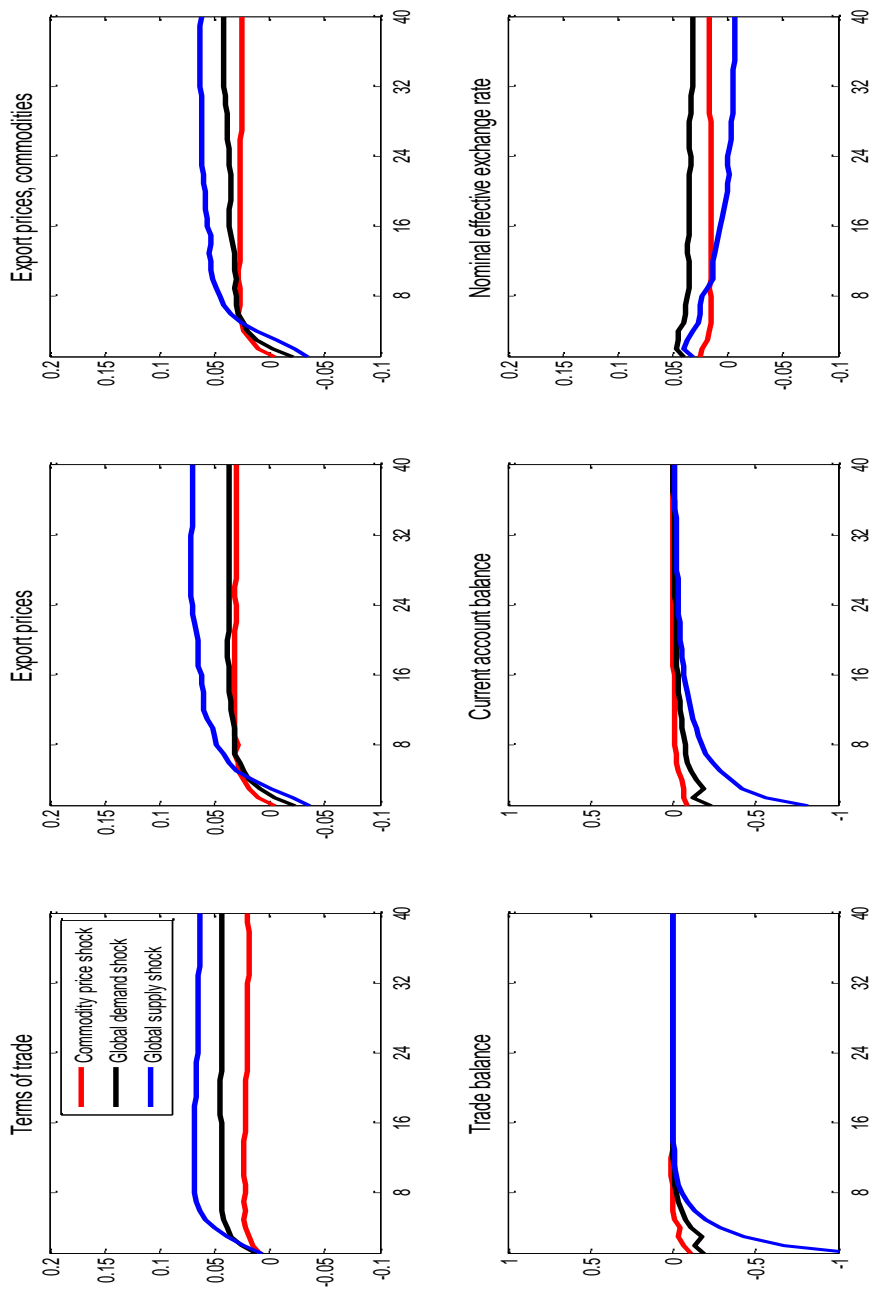
Figure 6: Impulse-response functions of Australian external balances and currency to global shocks. Median response.
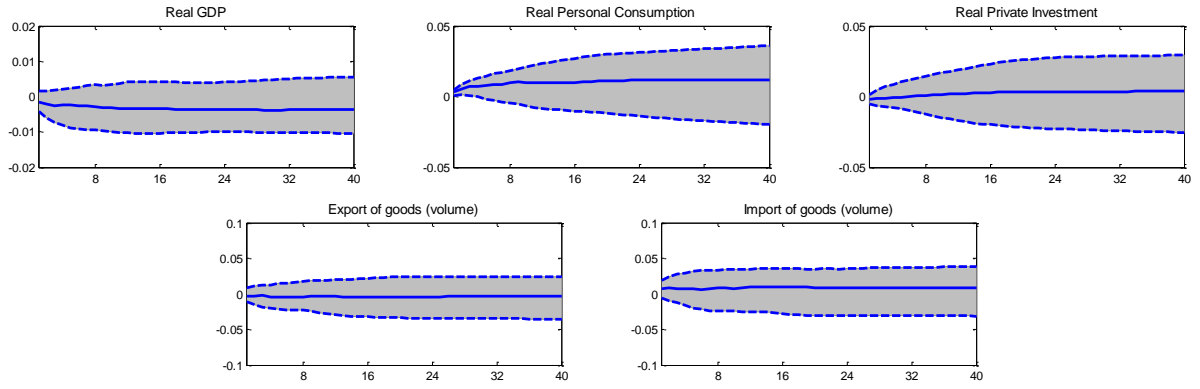
Panel A – Commodity price shock
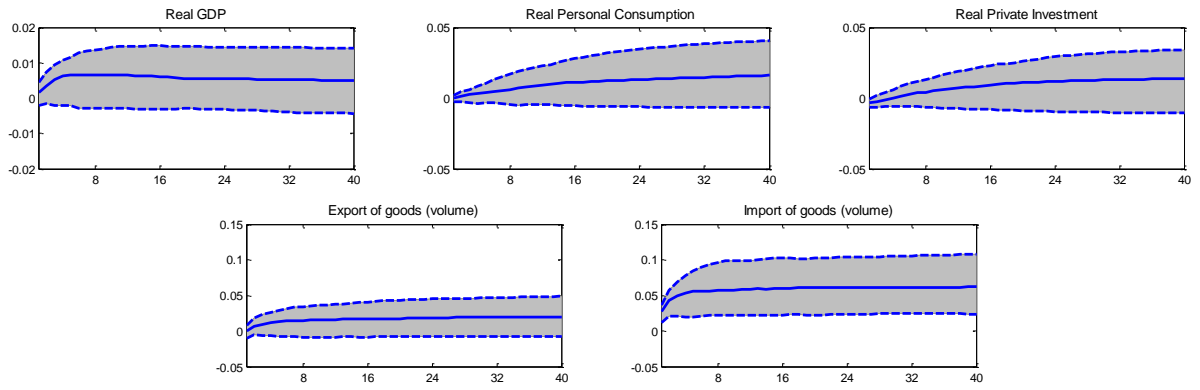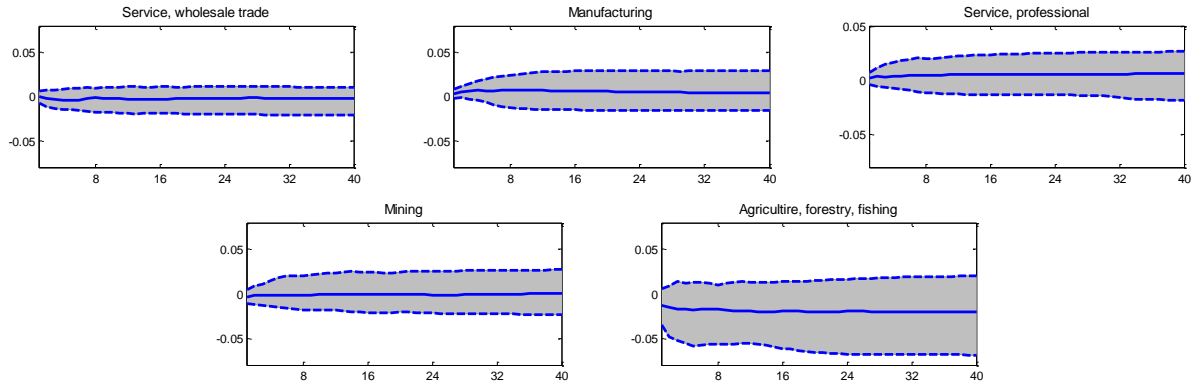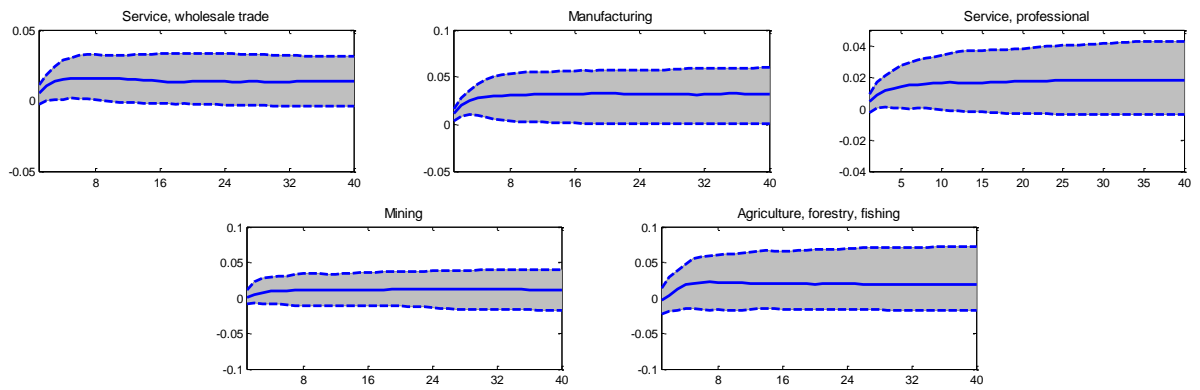

Panel B – Global demand
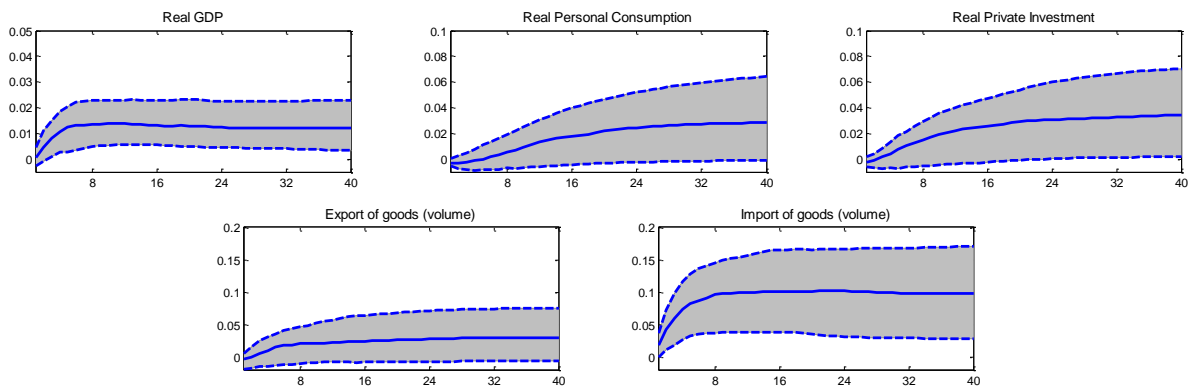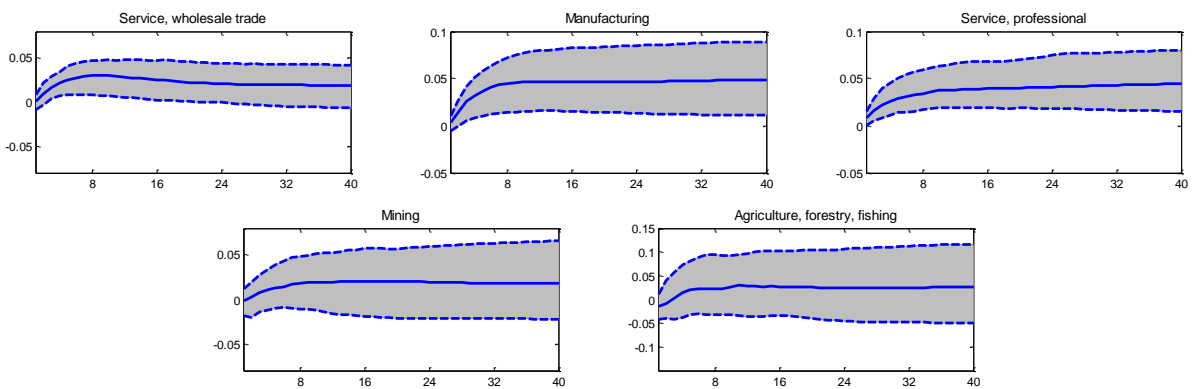

Panel C – Global supply shock

Figure 7: Impulse-response functions of GDP and spending aggregates to global shocks. 68% confidence bands.

Panel A – Commodity price shock

Panel B – Global demand shock

Panel C – Global supply shock

Figure 8: Impulse-response functions of GDP industries to global shocks. 68% confidence bands.

Appendix A   Estimation procedure

To estimate the model, I follow the estimation method used by Charnavoki and Dolado (2014) which proceed with likelihood estimation by multi-move Gibbs sampling by estimating the parameters of the DFM by alternatively sampling them from conditional posterior distributions. The factors in the transition equation (2) are modeled as a restricted structural VAR model. Since the model deals with a number of different dependent variables, it can be estimated as a system of seemingly unrelated regression equations (SURE). Specifically, it can be written as follows:

$$y_t = X_t \beta + v_t$$

Where y_t is a vector of dimension $K \times 1$ of dependent variables, $\beta = (\beta'_1, \beta'_2, \ldots, \beta'_K)'$ is a vector of parameters, X_t is a block-diagonal matrix with blocks x'_{kt} which contains the current and lagged values of the factors for the k-th variable and $v_t = (v_{1t}, v_{2t}, \ldots, v_{Kt})'$ is a vector of errors such that $v_t \sim N(0, \Sigma)$.

The restricted SVAR is estimated through a Bayesian method borrowed from Koop, Poirier and Tobias (2007). An independent normal-Wishart prior is used in the model which is largely employed in this type of model. A normal-Wishart can be written as

$$p(\beta, \Sigma^{-1}) \propto \phi(\beta | \underline{\beta}, \underline{R}) f_W(\Sigma^{-1} | \underline{T}, \underline{v})$$

Where $\phi(\cdot)$ and $f_W(\cdot)$ indicate Normal and Wishart probability density function, respectively. Accordingly, on the one hand, the conditional posterior distribution of restricted SVAR coefficients is

$$\beta | y, \Sigma^{-1} \sim N(\overline{\beta}, \overline{R})$$

With $\overline{R} = (\underline{R}^{-1} + \sum_{t=1}^{T} X_t' \Sigma^{-1} X_t)^{-1}$ and $\overline{\beta} = \overline{R}(\underline{R}^{-1}\underline{\beta} + \sum_{t=1}^{T} X_t' \Sigma^{-1} y_t)$.

While, on the other hand, the posterior for Σ^{-1} conditional on β is

$$\Sigma^{-1}|y, \beta \sim W(\overline{T}, \overline{v})$$

With $\overline{T} = (\underline{T}^{-1} + \sum_{t=1}^{T} (y_t - X_t\beta)(y_t - X_t\beta)')^{-1}$ and $\overline{v} = T + \underline{v}$. The prior is assumed to be uninformative, such that $\underline{R}^{-1} = \underline{v} = \underline{T}^{-1} = 0$. The Gibbs sampler employed sequentially draws from the normal $\phi(\beta y, \Sigma^{-1})$ and $f_W(\Sigma^{-1}|y, \beta)$ in order to approximate the posterior distribution in the model.

Appendix B Sign restrictions identification scheme

The alternative scheme to identify the structural model is based on imposing sign and bound restrictions on the impulse-response function following the procedure introduced by Rubio-Ramirez, Waggoner and Zha (2010). I follow the (iterative) computational strategy provided by Charnavoki and Dolado (2014). Suppose A_0 is the impact matrix obtained by Cholesky decomposing the reduced form variance-covariance matrix Ω and $\tilde{Q}$ is the identity matrix the global and regional block substituted by any rotational orthogonal 5×5 matrix with $\tilde{Q}\tilde{Q}' = I$. A new impact matrix is given by $\tilde{A}_0 = A_0\tilde{Q}$ where $\tilde{A}_0\tilde{A}_0' = \Omega$ and a number of structural models is obtained repeatedly drawing from the set of orthogonal rotational matrices. The procedure is articulated as follows:

- Cholesky decompose A_0^k of the posterior draw k of the reduced form variance-covariance matrix Ω^k.

- Suppose $X = QR$ where X is an independent standard normal 15×15, QR its decomposition with the diagonal of R, Q is a rotational matrix uniformly distributed. Substitute the global and regional diagonal block of $\tilde{Q}$ with Q.

- Compute $B_0^k = A_0^k\tilde{Q}$ and check if the model satisfies the sign and bounds constraints otherwise move to the next Gibbs iteration.

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Chapter 2

Forecasting business cycles in small open economies

Forecasting business cycles in small open economies[*]

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November 2018

Abstract

Do economic variables behaving as signals for incoming recessions exist? The answer to this question is provided through the estimation of an early warning system for a group of small open economies. Employing a number of indicators and dating domestic business cycles with a reliable method previously and largely used, both in-sample and out-of-sample results turn out to be satisfactory in terms of economic downturns predictions. Robustness exercises moreover show the dependence of the outcomes to the identification method of the economic phases and the presence of international spillover which drive recessions, both of them outside the control of national policy-makers.

Keywords: Recessions; small open economies; forecasting; business cycle identification.

JEL Classification: C32, E37, F41.

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1 Introduction

Are there economic indicators which behave as signals to warn about future downturns? The answer to this question has been largely provided in several contributions through the employment of leading indicators, whose literature was started by Burns and Mitchell (1938) and Burns and Mitchell (1946). Recently financial crisis testifies the need to take care of possible future downturns (developments) to stem economic slumps. The possibility to predict economic events is intended to be useful for policy makers at the time of policy decision and assessment as well as for private consumers or investors.

This paper estimates an early warning system to foresee economic downturns in the business cycle of a group of small open economies. In details, classical business cycle turning points are settle through a reliable method; afterward a state-dependent model is estimated, whose covariates are the main drivers of small open economies business cycle. Then the forecasting performance of the model is verified to test its feature of early warning system for economic contractions. Lastly the validity of the early warning system to a different business cycle identification is checked.

Twofold procedure is employed. In the first step business cycle phases are assessed through the methodology introduced by Bry and Broschan (1971) to detect peaks and troughs; subsequently a regime-conditional logit model is estimated, where lagged independent variables are loaded to build an early warning system for economic contractions.

The reason under the employment of the methodology by Bry and Broschan to date the business cycle relies on its over performing features in an ex-post characterization, as suggested by Harding and Pagan (2003); on the other hand, the motivations behind the logit approach are to ensure that the transition probability from one regime to the other is well define while including information from many variables.

This paper contributes to the literature on forecasting business cycle phases in several ways: (i) it builds an early warning system for recessions, extending the literature on early warning systems which are limited to currency and banking crisis so far; (ii) it focuses on main national economic indicators featuring the state of economy of small open

economies; (iii) it tests the functioning of the early warning system either in-sample and especially (and innovatively with respect to other early warning systems) out-of-sample; (iv) it checks and confirms that baseline results are driven by specific classifications of recessions, in particular with respect to the degree of their duration (Layton and Smith, 2007).

To carry the analysis I consider a group of four countries: Canada, Norway, Australia and France, which can all be intended as small open economies with respect to the United States. The data are collected at quarterly frequency and they span a period of 37 years ending in 2016. The dataset is composed by the series which best characterize the state of economy of small open economies (e.g. the trade balance and the terms of trade) and integrated with variables employed in the empirical framework of early warning systems (among the others, GDP growth and interest rate).

The results can be summarized as follows. First, the business cycle dating exercise shows either different timing of the turning points and the response to global economic events but it points out the common properties in terms of state persistence during specific periods. Then, concerning the opening question on the existence of economic indicators, the estimation of the model suggests the presence of variables able to signal future downturns; the reliability of the in-sample is tested through out-of-sample evaluations. Finally, controlling for a different dating approach for economic phases, arises evident facts connected to the method of crisis classification.

The paper develops as follows. Section 2 discusses the connection to the literature about the dating and forecasting business cycles and early warning system. Section 3 provides the description of the data and the empirical methodology to estimate the early warning system. Section 4 shows the results of the baseline model. Section 5 illustrates the results of the robustness checks. Section 6 concludes.

2   Literature review

The contributions concerning business cycle dynamics forecasts has been started by Burns and Mitchell (1938, 1946) in their analysis about turning points in the United States. The research field related to the issue of dating and forecasting the shape of the business cycles received its popularity in the last years. Since there are no authoritative dating of classical business cycles which can be regarded as the official reference cycle, except for the United States, whose turning points are officially defined by the Business Cycle Dating Committee of the National Bureau of Economic Research (NBER), the need to enlarge the knowledge of business cycle dynamics looks relevant.[1]

Several studies are focused on the role of different kind of data in driving business cycle along expansions and contractions. The analysis is almost everywhere performed with the use of composite index with different gauge, namely a short range gauge, a long range gauge and also an index of coincident business cycle indicators; the indices are used as explanatory variables in shaping future phases of the business cycle, in particular for the United States. The empirical set-up is built on Markov-switching models and sometimes on logit/probit models. Layton and Katsuura (2001) compare the performances of this two approach in forecasting the business cycle, while Moolman (2004) applies the two techniques to South African data using the yield spread as regressor documenting good outcomes for both models. Di Venuto and Layton (2005), followed by Layton and Smith (2007), employ a regime-dependent multinomial logit for the United States and Australia, respectively, to study to dependence of expansion and contraction to their respective duration; they document that the duration dependence still holds for contraction and expansion but by augmenting the model with new variables, other than composite indices, the results for the United States are different, particularly regarding the expansion period which show weaker dependence to its duration. The choice of detecting the business cycles in the present paper is mostly motivated by the lack of comparison among

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[1]Additional information about domestic business cycles are provided by the Economic Cycle of Research Institute (ECRI), which lists the turning points for 21 economies from 1948 to 2016 (1948-2015 for the Asian-Pacific and African countries).

similar and comparable countries in warning indicators responsiveness. Anas, Billio, Ferrara and Mazzi (2008), Darné and Ferrara (2011) and Billio, Casarin, Ravazzolo and van Dijk (2012) focus the analysis on identifying turning points for the Euro Area; Chauvet (1998), Chauvet and Piger (2008), Harding and Pagan (2002, 2006), Hamilton (2011) and Stock and Watson (2014) proceed to document the development of the business cycle in the United States; while other papers illustrate that predictions of variables at macro level can be gained via financial data (Estrella and Mishkin, 1998, Stock and Watson, 2003, Næs, Skjeltorp and Ødegaard, 2011 and Aastveit and Trovik, 2012); useful and effective prediction performance are also played by surveys data (Hansson, Jansson, and Löf, 2005, Abberger, 2007, Claveria, Pons and Ramos, 2007 and Martinsen, Ravazzolo and Wulfsberg, 2014). Logit and probit models received little attention to this aim. They have been applied for studies related to the United States; accordingly, the business cycle phases are obtained through the procedure by Bry and Broschan (for additional details, see Bry and Broschan, 1971 and Harding and Pagan, 2002), whose strategy has been borrowed to date the business cycle for the United States by NBER; the contraction and expansion phases are then employed as outcome to gather information on the existence of drivers of economic developments and slowdowns. There is a strong debate about forecasting performance of Bry and Broschan approach and Markov-switching models. Hamilton (2011) states there are difficulties in dealing with factors in dating turning points because of structural dynamic complexities of the factors themselves; previously Harding and Pagan (2003) show that ex post identification of the business cycle for United States is convenient via Bry and Broschan methodology with respect to Markov-switching models, while more recently Chauvet and Piger (2008) claim the efficiency of Markov-switching dynamic factor model (initially proposed by Chauvet, 1998) in detecting turning points.[2]

The choice of the explanatory variables relates to papers in the relevant literature of early warning systems and forecasting business cycles, with connection to small open economies framework. Covariates representing macroeconomic fundamental and monetary condition

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[2]For additional contributions using Markov-switching models, see, among the others, Filardo (1994) and Layton (1998), which employ leading indicators to predict business cycle phase shifts in the United States.

are common indicator employed to estimate early warning systems, especially related to systemic banking and currency crisis (Demirguc-Kunt and Detragiache, 1998; Kaminsky and Reinhart, 1999, Davis and Karim, 2008, Barrell, Davis, Karim and Liadze, 2010, Wong, Wong and Leung, 2010 Alessi and Detken, 2011, Babecký, Havránek, Matějů, Rusnák, Šmídková and Vašíček, 2013 and Caggiano, Calice and Leonida, 2014) where comparable and non-comparable economies are pooled together; while Aastveit, Jore and Ravazzolo (2016) estimate a Markov-switching dynamic factor model for identifying turning points in Norwegian business cycle performing a principal component analysis on a group of observed series strongly characterizing the profile of small open economy of Norway. Therefore, accordingly, the independent variables dataset is composed by a number of series at national level which can be directly or indirectly influenced by the policy makers; anyhow the dependence of the domestic business cycles to global occurrences will be verified through some robustness exercises.

3 Data description and empirical strategy

3.1 Data

The investigation is performed on a group of four countries: Canada, Norway, Australia and France[3], which can all be considered as small open economies with respect to the United States. To avoid to face sample selection issues, the analysis is extended to more countries (other commonly employed small open economies like the U.K. and New Zealand); the results are qualitatively consistent with the ones provided for the economies mentioned above even if the smaller availability of time observations narrows down its consideration to be included in the study; data at quarterly frequency are employed. The choice of data comes from a combination of relevant contributions on estimation of early

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[3]France is the only small open economy in the sample which cannot be considered as a commodity producer or exporter. The choice of including France is due to the possibility to control for countries other than commodity exporters (or producers), widening the analysis and checking the validity of the results for different types of small open economies.

warning systems and analysis of business cycles for small open economies. The dataset covers the period between 1980:I and 2016:IV. The variables comprised in the specification of the baseline model are one to four first principal components of four categories of covariates (respectively indicated as F1, F2, F3, F4):

- Macroeconomic indicators (F1): GDP growth, inflation and exchange rate depreciation. GDP growth is expected to influence the dynamic of the business cycle since it draws a picture of the economic development of the different industries; inflation is connected to the level of prices and, to some extent, to the investment decisions; the exchange rate plays a prominent role in determining the level of exports and imports, especially in this type of economies.

- Household conditions (F2): final consumption and unemployment rate. These two indicators connote the condition of the individuals as final consumers. The state of the economy is indeed severely contingent to the level of employment which is connected to the level of final private demand and consumption.

- Liquidity status (F3): narrow money (M1) or M3 and interest rate. The former detect the level of liquidity circulation among the consumers and can affect their purchasing power, while the latter is the useful items to manage the monetary policy.

- External activities variables (F4): change in the terms of trade and trade balance (as percentage of GDP). These two variables allow to identify to function of the external sector in predicting recessions; they represent important reference points of the economic activities for a small open economy.

Differently from prior papers on early warning systems, the data are not winsorized to

get the real picture of occurrences in the business cycles, even though including extreme observations may have an impact on the measurement error and, consequently, on empirical results. All the explanatory variables are lagged by one quarter, to see whether the model acts as an early warning system. It is worth saying that the employment of lagged variables enables to deal with potential endogeneity problems of the covariates.

3.2 Dating business cycles

The first step of the empirical strategy is connected to date the business cycles of small open economies. Business cycles are defined as "...a type of fluctuation found in the aggregate economic activity of nations that organize their work mainly in business enterprises: a cycle consists of expansions occurring at about the same time in many economic activities, followed by similarly general recessions, contractions, and revivals which merge into the expansion phase of the next cycle;..." (Burns and Mitchell, 1946, p.3) and its identification requires procedures suitable to respect its features. The methodology adopted is borrowed from Bry and Broschan (1971) and Harding and Pagan (2003): it consists of a non-parametric univariate approach to first detect potential peaks and potential troughs (Bry and Broschan, 1971) and then to select the actual turning points according to some criteria (Harding and Pagan, 2003). Formally peaks are defined when

$$P_t = (y_{t-2} \wedge y_{t-1}) < y_t > (y_{t+2} \wedge y_{t+1}) \tag{1}$$

while troughs

$$T_t = (y_{t-2} \wedge y_{t-1}) > y_t < (y_{t+2} \wedge y_{t+1}) \tag{2}$$

where y_t is GDP at quarterly frequency.

The rule of identifying a turning point inside a window of 5 quarters enables dynamic alternation of expansion and contraction periods as well as cycles and phases duration to validate the statement properties of Burns and Mitchell (1946). Following this approach eventually allows to coherently deal with the classical business cycle in the United States as defined by NBER. Borrowing from Aastveit, Jore and Ravazzolo (2016), the relationship between the business cycle (S_t) and the turning points can be interpreted as

$$S_t = S_{t-1}(1 - P_{t-1}) + (1 - S_{t-1})T_{t-1} \tag{3}$$

where S_t is equal to 1 in expansions and 0 in contractions; for instance, let's assume the economy is in recession, then $S_{t-1} = 0$; if $t - 1$ is a quarter when a trough takes place, $T_{t-1} = 0$ and so $S_t = 0$. S_t assumes value equal to 1 if there is a trough at $t - 1$, because in that case $T_{t-1} = 1$. S_t maintains its value equal 1 until a peak occurs.

3.3 Logit model

The core of the early warning system is built adopting the logit model, previously employed in other papers about early warning systems (Layton and Smith, 2007 and Di Venuto and Layton, 2005 for forecasting business cycles, Bussiere and Fratzscher, 2006 to foresee currency crisis and Caggiano, Calice and Leonida, 2014 in the context of systemic banking crisis). The purpose of the model is to convey an estimated probability of a specific phase of the business cycle to occur, where the probability is defined as a function of a vector of potential covariates by returning a predicted measure of the behavior of the business cycle. Specifically the logit model is constructed assuming that each economy can lay in one of the following $j + 1 = N$ states. Each economy has an associate probability to be in the state j that is given by

$$Pr(Y_{t=j}|X_t) = \frac{e^{\beta'_j X_t}}{1 + \sum_{l=1}^{J} e^{\beta'_l X_t}} \tag{4}$$

And the log-likelihood function to be maximized is

$$\ln(L) = \sum_{j=0}^{J} d_j \ln Pr(Y = j) \tag{5}$$

where X_t is the $k \times 1$ vector of independent variables and β is the vector of parameters to be estimated, while $d_j = 1$ denotes the case when economy is found to be in state j. The above shown log-likelihood function can be considered as a generalization of the log-likelihood function for the binomial logit model where each economy is allowed to lay in just two states, i.e. $Pr(Y_t > 1) = 0$; the log-odds ratio are instead returned as

$$\frac{Pr(Y_t = j)}{Pr(Y_t = 0)} = e^{\beta_j' X_t} \tag{6}$$

Where the vector of parameters β_j measures the effect of a change in the independent variables X_t on the probability for the country to enter a the state j. The state $j = 0$ is considered as the base outcome upon which providing identification for the logit model.

4 Results

4.1 Business cycle phases

The first results shown are about the identification of the business cycle phases subsequent to the turning points detection using through the procedure by Bry and Broschan (1971). As already specified in the previous paragraph, the dating process is performed on the data sample available at country level. Figure 1 plots the logarithm of national GDP along the sample periods spanned by each country. The results mostly show the presence of different dynamics inside the business cycles of the economies considered. Especially

for France, the economy appears to shift from one state to the other quite frequently; in the case of Australia, the business cycle exhibits a markedly stable persistence across the regimes. Generally speaking a lasting memory is common in all the countries during the 1990s consistently with information provided by ECRI and NBER which highlight approximatively no detection of turning points.

In Figure 2 the dynamic of unemployment rate is drawn; for Canada and Australia the recessions take place sistematically and correspondingly to unemployment rate rise; in Norway the economic slump in late 1980s is associated with a strong worsen in national employment.

Figure 3 shows the behavior of the trade balance (as % of GDP); it is immediate to note the distinction in terms of relevance of external activities with respesct to overall real activities in oil-exporting economies compared to importers of oil. Furthermore the way the trade balance performs next to recessions differs from country to country. It is worth additionally to say that world occurrences otherwise impact national business cycles via contagion processes. There are indeed evidences that the 2007-2008 financial crisis owns significant importance for the economic developments of the entire sample. Although the downturns connected to the event possess different durations across countries, the first quarter of the contraction period is located between the Q2:2008 and Q1:2009. Conversely the Dot-Com bubble seems to convey specific negative structural effects for the economic growth in Norway, France and Canada; in this case ECRI does not highlight any recession for France and Canada, even though some papers do (see e.g. Caggiano, Castelnuovo and Figueres, 2017, for Canadian business cycle). Table 1 summarizes these results by listing the identified turning points detected for each country. These differences across the economies, particular in terms of key variables, allow to monitor the effective function of the indicators in consistently predicting future contractions outbreaks and periods.

4.2 Model selection and in-sample results

The selection of the model for each country is based on a general-to-specific approach gradually loading the independent variables and comparing accordingly the results of two statistical measures. The fist indicator is the ROC curve introduced by Bergé and Jordá (2011) in the field of business cycle characterization initially employed in signal theory by Birdsall and Peterson (1953). The advantages of the ROC curve at the time of the quantification of fitting performance are, among the others:

- The independence to the underlying forecast loss function.

- The non-reliability to the occurrence frequency of specific states of the economy.

- The non-parametric estimation.

Actually the measure value of interest is the one obtained computing the area under the ROC (AUROC) curve; this helps to overcome the extra problem connected either to the choice of the cut-off threshold to recognize the recessions that are signaled and consequently to the trade-off between the recall ($specificity$) and the false positive ($1 - specificity$) rate. This is due to the correlation of the threshold with the degree of risk aversion (or tolerance) of the policymaker, hence the profile of his utility function. In the (unlikely) case of known utility function, the optimum threshold would be set so that the slope of the ROC curve equals the expected marginal rate of substitution between the net utilities of precise predictions. The values of the AUROC curve are contained in the interval [0.5,1], with a perfect informative model producing an AUROC=1, while a complete uninformative specification would generate an AUROC=0.5; hence a non-degenerated classifier (AUROC $\geq$ 0.5) is more informative as it produces an AUROC closer to 1.[4] The in-sample results are consistent along the whole sample of countries. The AUROC values are all above 0.80 for the models selected, meaning that the model adequately fits the data; the BIC results are in line with AUROC. Both measures support the model which comprises the entire set of factors as explanatory variables. The coefficients estimated are

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[4]The additional employed measure for model selection is BIC: the lower the value of the BIC, the better the fitting of the model.

generally statistically significant hence claiming the dependence of the business cycle to the indicators loaded (Table 2). Many commonalities are shared among the small open economies considered in terms of signals which warn the national economic system. In particular it is interesting to note that the factor extracted from the external activities is negatively correlated and with the probability of being in a recession quarter as well as, to a lesser extent, the factor connected to macroeconomic indicators. This may be intuitively understood since the share of the trade balance on overall GDP is not negligible given the economic structure of this type of economies. The stronger negative influence of the external variables on the probability of experiencing a recession testifies indeed the possibility the economic slump can be associated to the country reaction to the foreign environment condition, e.g. a flat or decreasing global demand. The monetary condition also signals for contractions; the results show that engaging excessive liquidity circulation triggers a surge in the interest rate which may lead to growth deceleration due to consumption and investment fall. The findings provide good reason to proceed with out-of-sample estimation and to check for effective early warning system function of the model.

4.3 Out-of-sample performance measures

The model chosen for each country is tested for its out-of-sample performance to verify its feature as early warning system. The out-of-sample results are based on one-step-ahead forecasting process with a rolling window approach. The length of the window is $m - 1$ where m is the quarter when the second recession of the sample begins.[5] Like for the in-sample analysis, two measures related to the out-of-sample results are provided. This choice is connected to the possibility to obtain a twofold forecasting skill valuation of the early warning system, one associated to the prediction of single events, the other to obtain an overall ability assessment.

Following previous papers estimating early warning systems (Demirg-Kunt and Detra-

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[5]It is necessary to include at least one recession in the rolling window to estimate the logit model. The length of the window for each country varies from 83 to 113 observations.

giache, 1998, Barrell, Davis, Karim and Liadze, 2010 and Caggiano, Calice and Leonida, 2014), the first measure is indicated as the in-sample threshold probability of recession to occur; the in-sample probability is estimated as the probability of a recession to occur inside the window whose last observation corresponds to a last quarter of expansion and a recession is said to be called if the estimated probability of recession at time m is higher than the in-sample probability of recession of the corresponding window. The second measure is the Brier score; the Brier score is a common indicator to evaluate the probability of an event to occur. The Brier score is computed as

$$ BS = \frac{1}{n} \sum_{t=1}^{n} (p_t - o_t)^2 \tag{7} $$

where n is the length of the sample, p_t is the estimated probability for an event to occur and o_t is a dummy equal to 1 if the event occurred and 0 otherwise. Understandably the lower the Brier score, the better the model out-of-sample performance.

As for in-sample results, the out-of-sample outcomes provide evidence of the existence of some properties common to the countries in the sample. First it is worth saying that almost all the recessions are correctly called when the in-sample probability is employed as acceptance threshold (Table 3-6); what is interesting to mention is based on the fact that the model is not always able to predict (or surely call) the 2008-2009 recession, confirming the uneasy predictability of this event; the forecasted probability is indeed smaller than the in-sample probability for Norway and France and slightly above the threshold for Canada. It is interesting to note that the recessions starting at t after the great financial crisis and that hence included this event in the window ending at $t-1$ are precisely predicted with probability values strongly greater than the in-sample likelihood. Secondly the Brier score values are all under 0.30 which can be intended as good finding for the reliability of the model as early warning system (Table 7).

5 Robustness analysis

The robustness of the results is now tested through the performance of some exercises. Essentially the checking for the external validity of the baseline model is twofold; first the resilience of the benchmark specification to different dating methods is verified; secondly the model is augmented with over national variables to control for possible external factors which may drive the outcomes. It is worth mentioning that, whether the early warning system performs better after these modifications or not, the variables are anyway not considered in the baseline model since they can not be under the control of national policymaker; it could be indeed an indication which may call for higher regional or global cooperation and integration to manage national economic difficulties.

5.1 Alternative dating

As first exercise, the robustness of the results is now verified through the consideration of a different dating approach. This appears necessary to check for the reliance of the baseline early warning system to the definition of the business cycle phases. The expansions and recessions as identified by the Economic Cycle Research Institute (ECRI) are therefore employed. Due to the convenience of using the same window length employed for the baseline estimation, this exercise is performed only for Canada and France. The frequency of ECRI business cycle statistics is monthly, hence the start and the end of the recession are set to the corresponding quarter; the model for each country is the one estimated in the baseline specification. Both in-sample and out-of-sample estimation do not report good results, at least compared with the previous findings. To the aim of overcoming the problem, it is easily noting that the occurrence of an economic contraction is more occasional with respect to one detected via the Bry and Broschan (1971) procedure. Specifically the recessions as dated by ECRI have an average duration of 5.75 quarters, while the ones identified in the baseline specification last 2.89 quarters, on average. This fact induces to attempt to consider the recession as not uniquely defined along its dura-

tion; this choice may indeed lead to a *recession duration bias*. Accordingly, a multinomial logit model with $j + 1 = 3$ states is thus employed to attest to this eventuality. Specifically the quarters of recession different from the first are classified separately from slump outbreak.

The new results are satisfying; the new classification and the extended model employed mostly lead to good outcomes, since all the crisis are indeed correctly signaled. Besides the computed brier score report smaller values compared to the ones estimated for the baseline model, meaning better forecasting performance. It is straightforwardly easy to state that the early warning system is somehow dependent to the classification of the dependent variable, namely business cycle. In particular it looks clearly that the recession is dependent to its duration.

5.2 The influence of external factors

The impact of global events can somehow and significantly influence the state of economy of a single country. This is why this eventuality is taken into consideration for this specific and additional robustness check. In particular, the baseline early warning system is augmented first with an oil price series and secondly with a dummy with takes value 1 if a recession in the U.S. takes place in the reference quarters and 0 otherwise, as indicated by NBER. The motivation behind these choices relies, on one hand on the fact that, except for France, all the countries in the sample can be identified as commodity-exporting economies, hence developments in commodity markets can definitely impact the domestic business cycle; on the other hand controlling for U.S. recessions helps to monitor predictable spillover effects spreading at global level. The new specifications are valuated with respect to their out-of-sample performances and significance of the coefficients in the in-sample estimation.

As for the previous check, it is easily noting that the model increases its forecasting skills when the early warning system is estimated including oil price and U.S. recessions (only

Norway's 2001 recession is not predicted according to the in-sample probability assessment method). It is worth precising that the coefficient of the U.S. recession dummy is positively correlated with the dependent variable, while oil price is negatively correlated with the probability of being in a recession quarter with respect to be in traquil times in oil-importing countries in the sample (Australia and France). This highlights the fact that other than internal determinants drive the state of economy, which may not be completely under the control of national policy makers.

6   Concluding remarks

This paper develops an early warning system for four small open economies, namely Canada, Norway, Australia and France employing quarterly data spanning the period between 1980:I and 2016:IV. Using the dating approach by Bry and Broschan (1971), a binomial logit model is built considering expansion and recession as the two states of reference. Four principal components estimated from relevant business cycle indicators for this type of economies. Through the employment of different performance measures, the covariates loaded (and lagged) in the baseline specification coveys their working efficiency in both in-sample and out-of-sample context to predict economic downturns. The validity of the results is anyway weakened when the business cycles are differently identified. In particular, consistently with previous contributions, the reliance of the recession to its duration turns out to matter and it is verified through the employment of multinomial logit model. Lastly the dependence of state of economy to external factors outside the control of domestic policy-makers is asserted, calling for higher supranational control and supervision.
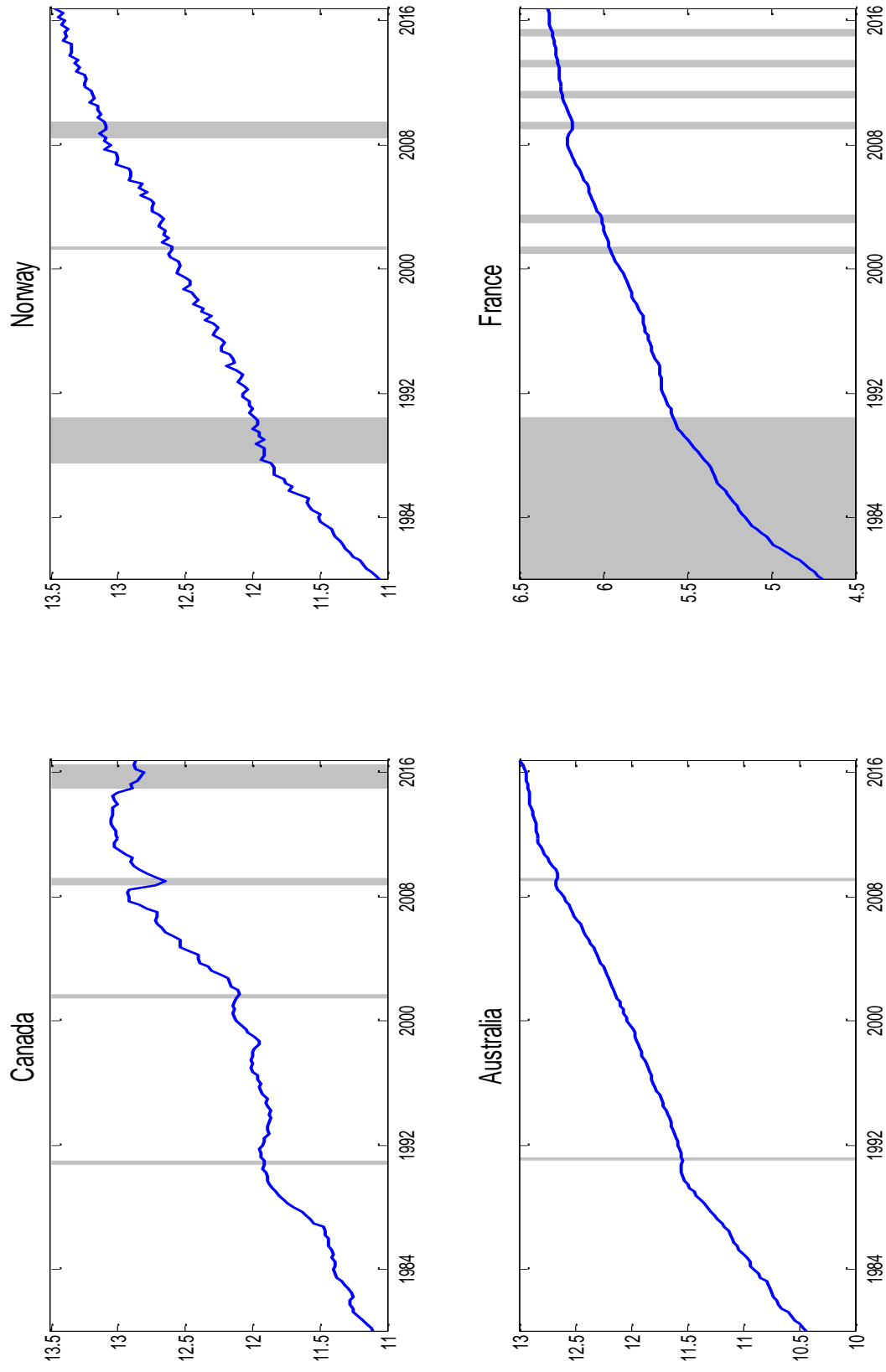
Figure 1: Log(GDP). Shaded areas indicate recessions identified through Bry and Broschan (1971) procedure.
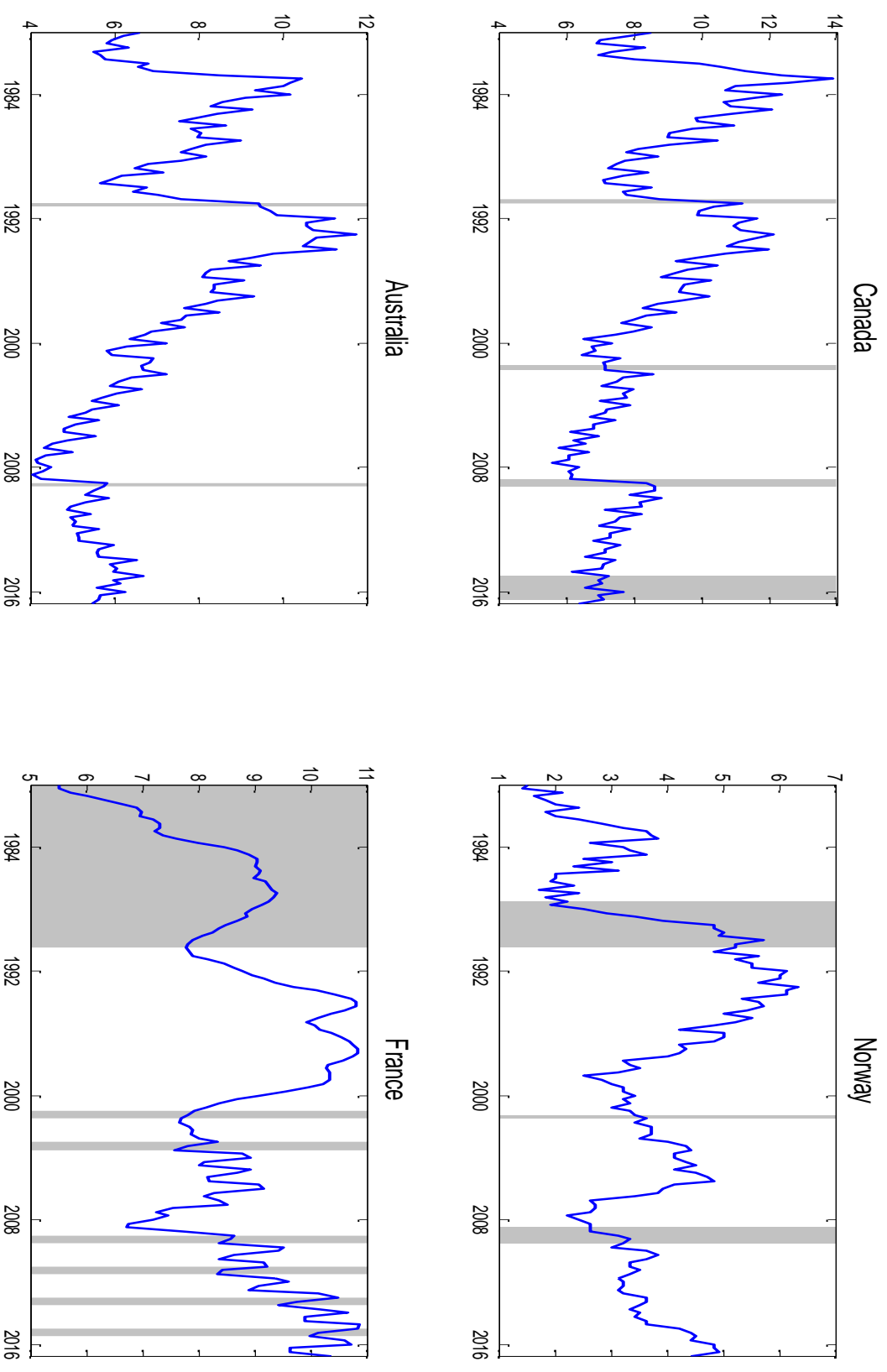
Figure 2: Unemployment rate. Shaded areas indicate recessions identified through Bry and Broschan (1971) procedure.
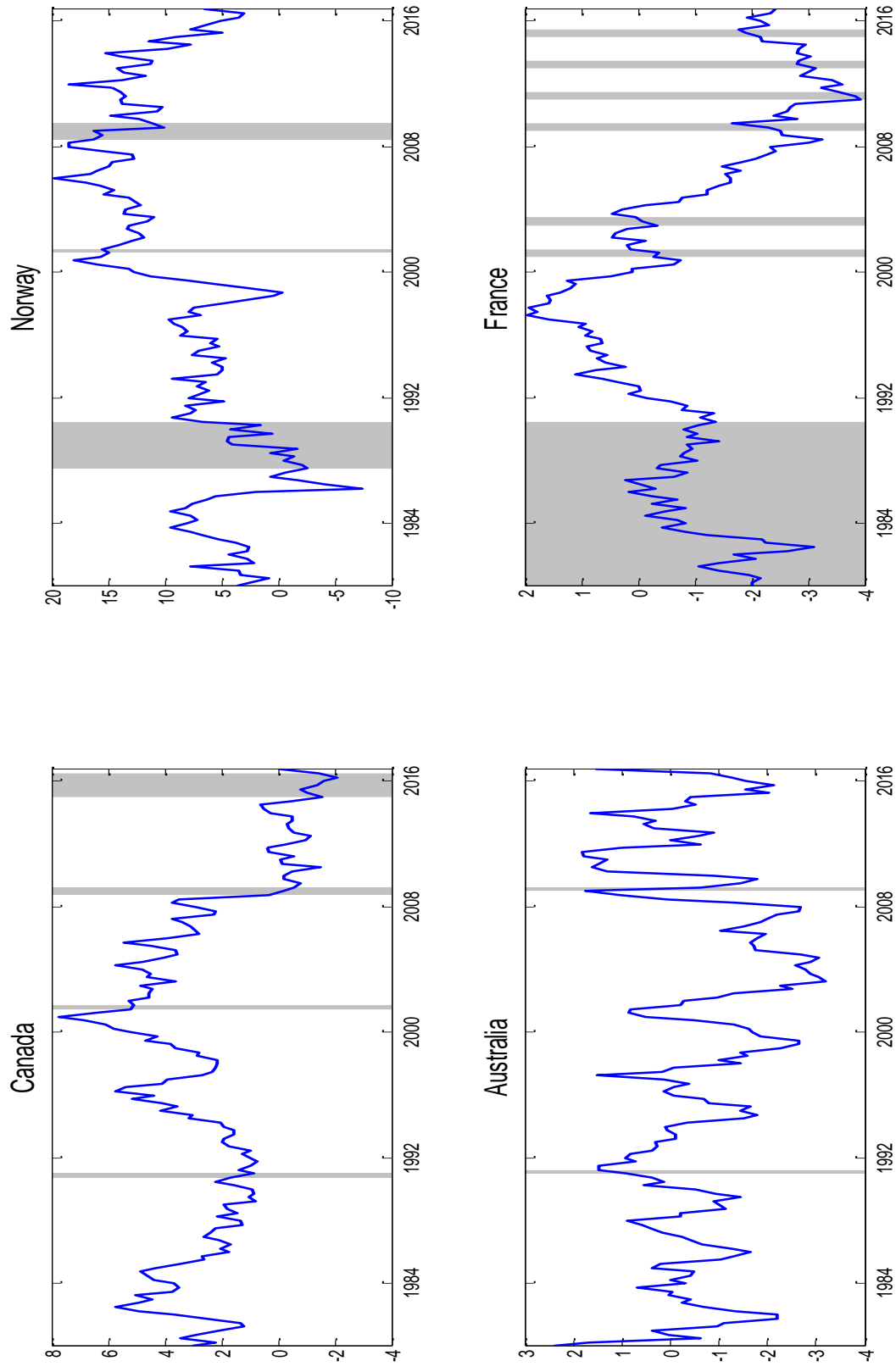
Figure 3: Trade balance (as % of GDP). Shaded areas indicate recessions identified through Bry and Broschan (1971) procedure.

Table 1: Turning points. Bry and Broschan (1971) dating approach.

Turning points	Canada	Norway	Australia	France
1980-1989				
Peak		1987Q2		
Trough		1990Q3		
1990-1999				
Peak	1990Q3		1990Q4	
Trough	1991Q1		1991Q2	1990Q3
2000-2001				
Peak	2001Q2	2001Q1		2000Q4
Trough	2001Q4	2001Q3		2001Q3
2002-2004				
Peak				2002Q4
Trough				2003Q3
2005-2009				
Peak	2008Q3	2008Q2	2008Q4	2008Q4
Trough	2009Q2	2009Q3	2009Q2	2009Q3
2010-2011				
Peak				2010Q4
Trough				2011Q3
2012-2013				
Peak				2012Q4
Trough				2013Q3
2014-2016				
Peak	2014Q4			2014Q4
Trough	2016Q3			2015Q3

Table 2: Area under the ROC curve for each country.

	Canada	Norway	Australia	France
AUROC	0.81	0.92	0.88	0.86

Notes: The AUROC curve values are the ones for the models with all the four factors loaded, which results the model with the best fitting.

Table 3: Coefficients of in-sample estimation

	F1	F2	F3	F4
Canada	-0.003***	-0.074*	0.063**	-0.009**
Norway	-0.012**	-0.039**	0.045*	-0.001***
Australia	-0.027**	-0.084**	0.019*	-0.032*
France	-0.029***	-0.038**	0.026**	-0.007*

Notes: ***, **, * indicate that the coefficients are significantly different from zero at the significance levels of 1%, 5% and 10%, respectively.

Table 4: Comparison between predicted and in-sample probability of recession for Canada

Recessions	Predicted probability	In-sample probability
2001Q3-2001Q4	0.013	0.012
2008Q4-2009Q2	0.052	0.024
2015Q1-2016Q3	0.407	0.012

Table 5: Comparison between predicted and in-sample probability of recession for Norway

Recessions	Predicted probability	In-sample probability
2001Q2-2001Q3	0.033	0.011
2008Q3-2009Q3	0.006	0.02

Table 6: Comparison between predicted and in-sample probability of recession for Australia

Recessions	Predicted probability	In-sample probability
2009Q1-2009Q2	0.630	0.009

Table 7: Comparison between predicted and in-sample probability of recession for France

Recessions	Predicted probability	In-sample probability
2001Q1-2001Q3	0.429	0.012
2003Q1-2003Q3	0.150	0.024
2009Q1-2009Q3	0.000	0.036
2011Q1-2011Q3	0.137	0.048
2013Q1-2013Q3	0.473	0.048
2015Q1-2015Q3	0.375	0.060

Table 8: Brier score estimation for each country

	Canada	Norway	Australia	France
Brier score	0.11	0.06	0.18	0.14

Notes: the Brier score is estimated for the entire sample period. Understandbly, the lower the Brier score the better the out-of-sample estimation.

Table 9: Robustness analysis. Multinomial logit model coefficients estimation.

	F1 j=1	F1 j=2	F2 j=1	F2 j=2	F3 j=1	F3 j=2	F4 j=1	F4 j=2
Canada	-0.038**	-0.045*	-0.011**	-0.030**	0.015*	0.023***	-0.013***	-0.006
Norway	-0.008**	-0.022**	-0.021	-0.024**	0.001*	0.037*	-0.020**	-0.003*
Australia	-0.014*	-0.019***	-0.005**	-0.013	0.029	-0.026**	-0.014*	0.009
France	-0.052***	-0.078***	-0.009**	-0.046***	-0.068*	-0.004**	-0.005*	-0.002

Notes: ***, **, * indicate that the coefficients are significantly different from zero at the significance levels of 1%, 5% and 10%, respectively.

Table 10: Robustness analysis. Comparison between predicted and in-sample probability of recession for Canada. ECRI turning points.

Recessions	Predicted probability	In-sample probability
2008Q1-2009Q3	0.334	0.012

Table 11: Robustness analysis. Comparison between predicted and in-sample probability of recession for France. ECRI turning points.

Recessions	Predicted probability	In-sample probability
2002Q3-2003Q2	0.301	0.024
2008Q1-2009Q1	0.362	0.036
2011Q2-2012Q4	0.374	0.048

Table 12: Robustness analysis. Brier score estimation for Canada and France. ECRI dating.

	Canada	France
Brier score	0.08	0.009

Notes: the Brier score is estimated for the entire sample period. Understandbly, the lower the Brier score the better the out-of-sample estimation.

Table 13: Robustness analysis. In-sample coefficients estimation.

	Oil price	U.S. recession dummy
Canada	0.025***	0.114***
Norway	0.071**	0.022*
Australia	-0.005	0.009*
France	-0.031	0.076***

Notes: ***, **, * indicate that the coefficients are significantly different from zero at the significance levels of 1%, 5% and 10%, respectively.

Table 14: Robustness analysis. Comparison between predicted and in-sample probability of recession for Canada

Recessions	Predicted probability (oil price)	Predicted probability (U.S. recession dummy)	In-sample probability
2001Q3-2001Q4	0.121	0.146	0.012
2008Q4-2009Q2	0.137	0.299	0.024
2015Q1-2016Q3	0.316	0.384	0.012

Table 15: Robustness analysis. Comparison between predicted and in-sample probability of recession for Norway

Recessions	Predicted probability (oil price)	Predicted probability (U.S. recession dummy)	In-sample probability
2001Q2-2001Q3	0.355	0.002	0.011
2008Q3-2009Q3	0.247	0.041	0.023

Table 16: Robustness analysis. Comparison between predicted and in-sample probability of recession for Australia

Recessions	Predicted probability (oil price)	Predicted probability (U.S. recession dummy)	In-sample probability
2009Q1-2009Q2	0.022	0.017	0.009

Table 17: Comparison between predicted and in-sample probability of recession for France

Recessions	Predicted probability (oil price)	Predicted probability (U.S. recession dummy)	In-sample probability
2001Q1-2001Q3	0.121	0.146	0.012
2003Q1-2003Q3	0.137	0.299	0.024
2009Q1-2009Q3	0.316	0.384	0.012
2011Q1-2011Q3	0.137	0.299	0.024
2013Q1-2013Q3	0.034	0.249	0.024
2015Q1-2015Q3	0.181	0.225	0.060

Table 18: Robustness analysis. Brier score estimation for each country.

	Canada	Norway	Australia	France
Brier score (oil price)	0.07	0.04	0.23	0.15
Brier score (U.S. recession dummy)	0.05	0.17	0.31	0.10

Notes: the Brier score is estimated for the entire sample period. Understandbly, the lower the Brier score the better the out-of-sample estimation.

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Chapter 3

Small open economies and inflation forecasts with global and regional components

Small open economies and inflation forecasts with global and regional components

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Abstract

The paper considers the dynamics of inflation in a group of small open economies. Through the Bayesian estimation of a set of linear models it is shown that the predictability of national inflation is strictly dependent to its degree of control of the persistence. Furthermore and more relevantly the evidence suggests that global inflation does not produce better forecasting performance for national price level with data employed monthly frequency. This issue is therefore faced including regional inflation for all the countries in the sample; the results obtained highlight the importance of sub-global business cycle in foreseeing national price level change either at point and density level.

Keywords: Inflation; small open economies; regional inflation; point and density forecast.

JEL Classification: C32, E31, E37.

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1 Introduction

Ciccarelli and Mojon (2010) can be considered as the first authors deeply testifying the role of international common inflation dynamics and its function in capturing national current and future price level developments. Price inflation forecasts indeed rely among the hottest topics for policymakers to be dealt with. The price movement and, to the same extent, its predictions indeed denote one of the main drivers for economic stability and they are the primary objectives for most of the national (e.g. the Federal Reserve Bank) and supranational (e.g. the European Central Bank) central banks. The issue of time-varying economic conditions and individual expectations makes the existing procedures sometimes obsolete and sometimes incomplete; because of this the probability of price-related economic issues might progressively increase[1] and the results could be not useful enough to promptly face incoming negative economic conditions if obtained through unique models estimations and single predicting performance. On the one hand, the former stands for reliability of basic standard models whose validity may be compared to additional different schemes, on the other hand, the latter states that entrusting the whole forecast framework to a unique forecasting ability measurement may come out empirically restrictive.

The purpose of this paper is to move along these two directions, extending meanwhile the evidence provided by Ciccarelli and Mojon (2010). In fact this work aims at widening the set of linear models to see whether new specifications increase the forecasting performances for inflation; this procedure is twofold in the sense that the predictive ability of the estimated models are compared under the lenses of two different levels of empirical predicting achievements, namely the point and the density forecast. The analysis is performed on the monthly inflation rate for a group of countries commonly identified as small open economies; the choice for this sample is due to the fact that the strong international market orientation of this type of economies can be exploited as motivation to outline

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[1]It is easy to think about the level of the federal funds rate which is near the zero lower bound, which generates ineffective monetary policy. More suitable models to forecast future price dynamics may have decreased the unexpected consequences of the financial crisis and the subsequent sovereign debt crisis.

the role of specific external spillovers in leading predictions improvements and identified as exogenous dimensions to be included in the benchmark specification. The stimulus to carry out such comprehensive exercises relies on their current absence in the related literature. In particular three developed countries are employed, i.e. Canada, Norway and the U.K.; the reliability of their inflation control is an additional factor motivating the selection of these countries; the monthly national inflation rate is forecasted at different future monthly horizons. Accounting for predictions on various months-ahead basis is justified by the different settled inflation targets which are commonly to be respected in the medium term.[2] The forecasting[3] performance is analyzed subsequently to the recursively estimation of standard autoregressive (AR) models and a number of generalized Bayesian AR models augmented with further lags and exogenous variables (BARX) . In particular, I consider additional variables to be added in the baseline model, namely the global and the regional inflation; moreover two different prior distributions of the variables upon which the Bayesian estimation is executed are assumed. The choice for the Bayesian estimation instead of the classical statistical methods relies mainly on its ability to work well in highly uncertain economic environments making the estimation more dependable. The outcomes show that at both point and density level, the forecast of specific BVAR models overperform with respect to the baseline (AR) and other BVAR models, but some results substantially differ from the ones by Ciccarelli and Mojon (2010). It is worth citing that the evidence is robust for the three countries in the sample and for both prior distributions. Especially relevant and necessary appears the inclusion of regional next to global inflation to increase the forecasting efficiency of the linear models assessed.

The paper proceeds as follows. Section 2 provides a related literature review. Section 3 presents the empirical framework and the data employed in the analysis. Section 4 shows the results. Section 5 concludes.

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[2]The inflation target rate in the European Union is defined on yearly basis. In this paper the inflation rate is employed at monthly frequency; this is however a good exercise to analysis forecast ability in the medium term.

[3]The forecasts are obtained iteratively. For completeness they could be obtained directly too, in order to test the consistency of the results.

2 Literature review

International synchronization is widely taking place among the contributions related to output dynamic, mainly connected to national real activities. Indeed, particularly since, among the others, Kose, Otrok and Whiteman (2003)[4] , the number of authors focused on global business cycles comovements has hugely increased. Recently, next to the world dimension of variables, regional developments have been included in new specifications for business cycle identification (Crucini, Kose and Otrok, 2011 and Mumtaz, Simonelli and Surico, 2011), stabilization (Thorsrud, 2013 and Aastveit, Bjørnland and Thorsrud, 2016) and output prediction (Bjørnland, Ravazzolo and Thorsrud, 2017).

Fewer attentions have been put on price inflation and its international connotation. A revisited version of the Phillips curve has been provided by Stock and Watson (1999) to test its stability and increasing performance in terms of forecasting ability. Ciccarelli and Mojon (2010) highlight the function of the global inflation in explaining national price levels with an out-of-sample exercise, while earlier Rogoff (2003) and Levin and Piger (2003) claim a common disinflationary trend on global scale and absence of strong persistence in national inflation series[5] , respectively. Mumtaz and Surico (2012) highlight the existence of country-specific factors, since inflation fluctuations are associated to the behavior of policy makers at the time of carrying out national economic policies. In-sample performances of global output in tracking national price levels have been described by Milani (2010), claiming that world output dynamic has a significant correlation with domestic demand and, indirectly, it affects countries inflation. Empirical evidence of the globalization of inflation has been provided by Bianchi and Civelli (2015), which show that inflation dependence to the world fluctuations is not time-varying but is positively related to trade openness. New works concerning the Phillips curve specified for open economies verify the existence of world dynamic for price inflation (Castelnuovo, 2010) and its role in

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[4]See also Forni and Reichlin (2001) and their study about regional policy integration for Europe and the U.S.; Canova, Ciccarelli and Ortega (2007) detect the importance of world indicator over the national variables in business cycle synchronization.

[5]See also Angeloni, Aucremanne and Ciccarelli (2006) and their results about whether EMU constitution affected price setting and inflation persistence in Euro countries.

shaping forecasters expectations in the U.S.; while and Eickmeier and Pijnenburg (2012) contribute to the literature on traditional Phillips curve underlining that the common component of changes in unit labour costs is not negligible to shape inflation dynamic for OECD countries; previously Atkenson and Ohanian (2001) employed the Phillips curve to forecast inflation.

Inflation forecasting exercises have been carried out through the employment of leading indicators (Banerjee, Marcellino and Masten, 2005, for Euro Area and Banerjee and Marcellino, 2006, for the U.S.), Bayesian model averaging (Groen, Paap and Ravazzolo, 2010) and heuristic optimization (Kapetanios, Marcellino and Papailias, 2016). On the same line of Ciccarelli and Mojon (2010), Pinchiera and Gatty (2016) test the power of regional and global inflation in predicting Chilean inflation.

3 Empirical set-up

In this section, the models used to forecast the measure of inflation provided in the data description section are introduced. As already anticipated, both univariate and multivariate models for monthly month-on-month inflation rate are employed with the addition of exogenous variables relative to regional and global inflation. The benchmark specification is a standard AR model with lag $p = 1$; the estimation is performed with Bayesian methods.

VARs are considered as multivariate models. The VARs have two different specifications: the one related to national and global inflation, the second which additionally includes regional inflation. The reason behind the first specification is to test the robustness of the results by Ciccarelli and Mojon (2010) connected to a different frequency of the inflation rate. The exogenous variables, as well as the endogenous, (exogenous with respect to the national inflation, i.e. the regional and global one) enter in the model with 2 lags each. The explanation connected to this choice is associated to the consistent necessity to control for persistence of the inflation also at supranational level. Specifically they are

$$\begin{pmatrix} \pi_t^* \\ \pi_t^W \end{pmatrix} = \sum_{c=1}^{p} \Phi_c \begin{pmatrix} \pi_{t-c}^{**} & \pi_{t-c}^{*W} \\ \pi_{t-c}^{W*} & \pi_{t-c}^{WW} \end{pmatrix} + u_t \tag{1}$$

and

$$\begin{pmatrix} \pi_t^* \\ \pi_t^W \\ \pi_t^R \end{pmatrix} = \sum_{c=1}^{p} \Phi_c \begin{pmatrix} \pi_{t-c}^{**} & \pi_{t-c}^{*W} & \pi_{t-c}^{*R} \\ \pi_{t-c}^{W*} & \pi_{t-c}^{WW} & \pi_{t-c}^{WR} \\ \pi_{t-c}^{R*} & \pi_{t-c}^{RW} & \pi_{t-c}^{RR} \end{pmatrix} + u_t \tag{2}$$

where π_t^*, π_t^W and π_t^R stands for national, global and regional inflation, respectively; Φ_c represents a matrix of coefficient of every covariates and lag multiplied by the general matrix of the lagged dependent variable; u_t indicates the error.

The univariate models are two: an AR(1) (baseline) and an AR(2). The second specification is estimated to be consistent with the multivariate models in terms of control of inflation persistence. In formulas they are represented as

$$\pi_t^* = \beta \pi_{t-1}^* + \epsilon_t \tag{3}$$

and

$$\pi_t^* = \sum_{c=1}^{p} \beta_c \pi_{t-c}^* + \epsilon_t \tag{4}$$

The data are collected from the database of the FRED (Federal Reserve Bank of St. Louis). They refer to three countries, namely Norway, Canada and the U.K. The inflation rate is computed as the month-on-month percentage change of the monthly national consumer price index. The same logic is applied to compute the regional and global inflation rate; the first refer to the Euro Area inflation rate for Norway and the UK and

to the U.S. inflation rate for Canada; for the second one the OECD countries consumer price index percentage change is considered.

4 Results

The results of the empirical analysis are now reported; the results are based on h-step-ahead forecasting process with a rolling window of 80 observations for each country. Earlier than the outcomes of the central analysis, some preliminary evidence related to the inflation properties are highlithed. In Figure 1 the monthly inflation rate for each country along the sample period is plotted. It is easily noting that, in line with Boivin and Giannoni (2009), the inflation rates decrease since the beginning of the 1980s. The main spikes are depicted during the main economic occurrences either at national and global level. For Norway, the main increase in inflation rate volatility takes place in correspondence of the economic recession of 2003[6]; the story is similar for the UK where inflation stability decreased at the time of the exchange rate mechanism crisis in 1993-1994. Canadian price level shows higher instability during the downturns at the beginning of 1990s.[7]

Statistical properties and evidence continue taking into consideration the variance explained by global and regional inflation; the share of the national inflation variance captured by regional price level is computed as

$$expl.variance_m = \alpha_m^2 var(\pi_s^R)/var(\pi_m^*) \tag{5}$$

and the for world inflation as

$$expl.variance_m = \alpha_m^2 var(\pi_s^W)/var(\pi_m^*) \tag{6}$$

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[6]There is no authorative dating for Norwegian business cycle. The evidence about the recession of 2003 is drawn from Aastveit, Jore and Ravazzolo (2016).

[7]Charnavoki and Dolado (2014) highlight the presence of a structural break in 1990 in the Canadian economy.

where α is the window-on-window OLS estimate when national inflation is regressed on regional or world inflation and the subscript m indicates the m-th window. The window-on-window explained variance is drawn in Figure 2. The figure shows that the regional inflation plays a key role in capturing domestic infaltion rate deviations. Its importance is particularly stressed during the first part of the sample period for the U.K. and, to a lesser extent, for Norway, while for Canada the presence of U.S. change in consumer price index is costantly significant. For completeness the cumulated explained variance is also computed, specifically

$$cum.expl.variance_m = \sum_{l=1}^{m} expl.variance_l \tag{7}$$

either for global and regional inflation.

Figure 3 testifies the increasing burden carried by regional inflation along the sample period. Especially for Cananda, neglacting regional inflation rate would connote statistical and economic inaccuracy. For the U.K. and Norway the sotry is diffent but sub-global inflation rate still represents a good proportion of the domestic cumulated explained variance.

These graphs represent the preliminary analysis prior to the further development of the present paper. Due to the graphical outcomes provided, the analysis was worth proceeding.

The estimation of the forecasting models is performed through the employment of two different prior to define a priori information on the matrix of coefficients and on the variance-covariance matrix; this choice has been made to test the robustness of the results in term of distribution of the data. The prior chosen are, specifically, the conjugate Normal-Wishart prior and the Minnesota prior for which the Gibbs sampling algorithm is used; all the models are estimated thorugh the Bayesian Markov chain Monte Carlo

(MCMC) methods.[8]

Before analyzing the out-of-sample findings, it is apposite to mention the fact that the in-sample evidence provides positive and statistically significant coefficients for the three countries; this confirms the relationship of national inflation to international price movements. As a matter of fact, it is appropriate to keep on with the out-of-sample analysis. The results concerning the out-of-sample performance of the models are analyzed under the lense of point and density forecasting for time horizons $h = 3, 6, 9, 12$. The former is evaluated through the computation of the root mean square errors (RMSEs), namely

$$RMSE = \sqrt{\frac{1}{T - M} \sum_{t=M}^{T-1} (\hat{\pi}^*_{t+h|t} - \pi^*_{t|t})} \tag{8}$$

where T represents the length of the sample, M is the dimension of each window and $\hat{\pi}^*_{t+h|t}$ are the inflation forecasts.

Density forecast is assessed estimating the log predictive score (LPS), which is seen as the widest employed measure for density predictability[9], specifically

$$LPS_t(\pi^*_{t+h}) = ln(f(\pi^*_{t+h}|I_t)) \tag{9}$$

where $f(\pi^*_{t+h}|I_t)$ is the predictive density for π^*_{t+h} computed conditionally on the information set available up to time t. The results about these two measures are reported to simplify and make their comprehension immediate. Indeed RMSEs and average log scores for the baseline AR(1) model are reported, while for the other models the ratios of each model's RMSE to the AR(1) model and the differences score relative to the baseline AR(1) model are displayed, where a ratio smaller than 1 and a positive difference establish better performance in terms of point and density, respectively, with respect to the benchmark AR(1).

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[8]The models are estimated following Koop and Korobilis (2010).
[9]For additional details see Geweke and Amisano (2010).

The results are provided in Table 1-12; it is worth pointing out that the Diebold-Mariano test is used to compare the forecasts. Table 1-6 and Table 7-12 show consistency between the evidence connected to both types of forecast, respectively point and density forecast; evidence and comparison connected to the two different prior specifications (i.e. Normal-Wishart and Minnesota) can be observed and make in Table 1-3 and Table 6-9 for Normal-Wishart prior and Table 4-6 and Table 10-12 for Minnesota prior; outcomes strictly focused on forecasting performance at country level are available comparing Table 1-4-7-10 for Canada, Table 2-5-8-11 for the UK and Table 3-6-9-12 for Norway. The main findings which worth focusing on are essentially two. First it is easily noting that, in line with several contributions, persistence is a feature that strongly characterizes the inflation process and that controlling for inflation memory helps to obtain gains in the forecast accuracy. In fact adding a lag to the baseline AR(1) produces better point and density predictions at every time horizon considered. The second and more relevant finding concerns the role of global and regional inflation and their impact on national inflation predictability integrated through the VAR models. Following Ciccarelli and Mojon (2010), global inflation dynamic is tested at monthly frequency this time. Augmenting the baseline model with 2 lags of world inflation however does not produce better results, either compared with the AR(1) and the AR(2) model. This outcome may generate some doubts about the global price dimension raising the reflection about the possible existence of some dependence of world inflation to data frequency. The step forward carried out concerns the additional inclusion of the exogenous variable depicting regional inflation. Here the improvements look really significant. The new variable indeed helps to reap the benefits that foreign variables, specifically regional variables, are widely tested to produce in terms of statistic gains (Aastveit, Bjørnland and Thorsrud, 2016 and Bjørnland, Ravazzolo and Thorsrud, 2017). Regional developments are therefore not negligible when dealing with international business cycle and in particular, among the others, price predictions. It seems reasonable to infer, on one hand, that lower data frequency may limit forecast ability of global variables; on the other hand, the reducing the scale of the international indicators seems like a smart move to positively influence the predictive skills of

aggregate inflation.

5   Concluding remarks

The paper valuates the influence of foreign inflation in the context of forecasting national inflation for the sample spanning the period comprised between 1975 and 2016. The data are collected at monthly frequency. Three small open economies are considered, i.e. Norway, the U.K. and Canada. Though a Bayesian procedure a group of univariate and multivariate models is estimated. The results concerning point and density forecasting suggest two main conclusions; first the inflation and the degree of its predictability are consistently dependent to the extent to which the persistence is controlled; secondly there is a possible connection dependence between the data frequency and the forecast ability of the international inflation; indeed increasing the data frequency with respect to the seminal work by Ciccarelli and Mojon (2010), it is estimated a decreased predictive performance of the model when global inflation is included next to the national price level. The further development of the present paper is to attempt to lower the international dimension of the foreign inflation by considering regional inflation together with national and global price change. The outcomes provide evidence of better results in terms of point and density forecast in this last case, in line with previous contributions which highlight the increasing importance of regional factors for country's indicators.
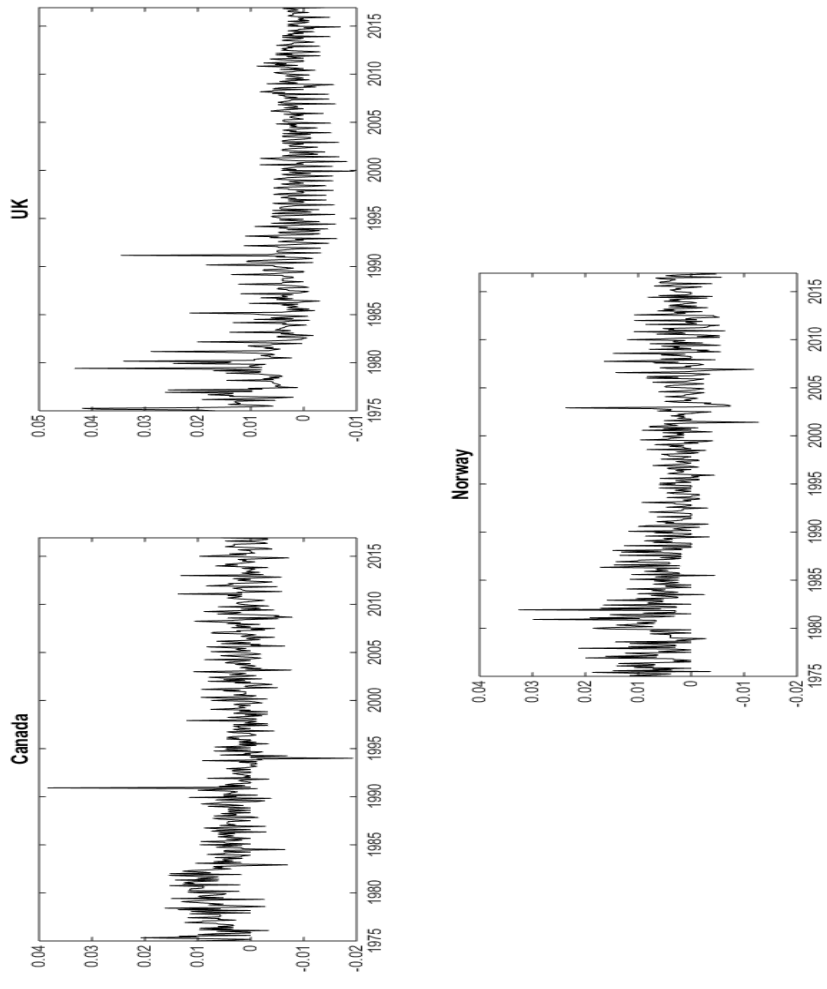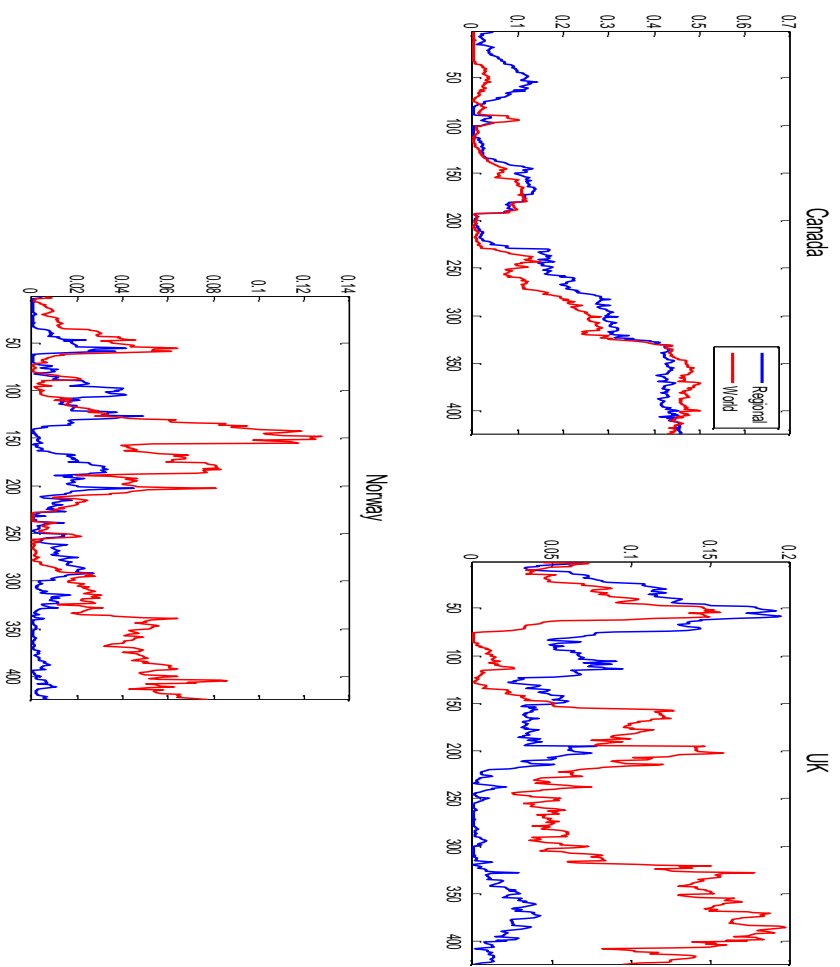
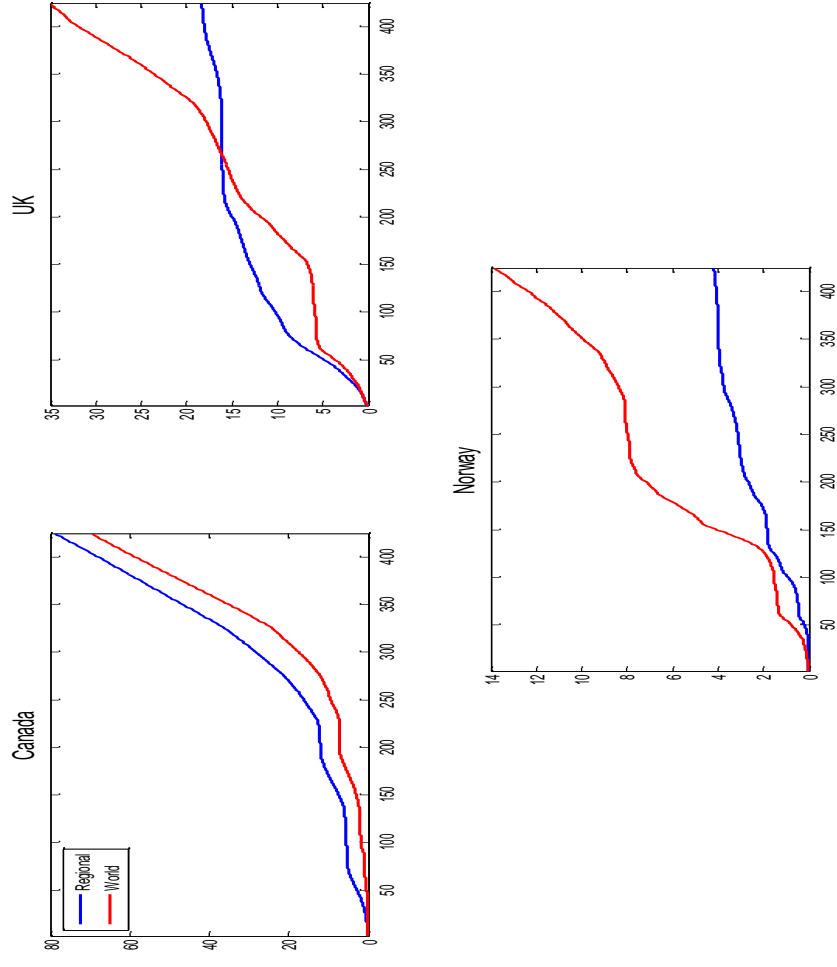Figure 1: Inflation rate.

Figure 2: Explained variance.

Figure 3: Cumulated explained variance.

Table 1: Point forecasts for Canada. Normal-Wishart prior

h	3	6	9	12
$BAR(1)$	15.287	19.327	15.453	17.814
$BAR(2)$	0.954***	0.935***	0.889***	0.904***
$BVAR_W$	1.027	1.104	1.031	1.096
$BVAR_{W,R}$	0.779**	0.894***	0.767**	0.848**

Notes:
- The table reports (i) the value of the RMSEs for the benchmark model BAR(1) and (ii) the ratio between the RMSE of each of the other models and the RMSE of the AR benchmark.
- ***, ** and * indicate that the RMSE ratios are significantly different from 1 at the significance levels of 1%, 5% and 10%, respectively, according to the Diebold-Mariano t-statistic test for equal RMSEs.

Table 2: Point forecasts for the UK. Normal-Wishart prior

h	3	6	9	12
$BAR(1)$	9.793	11.571	10.240	11.382
$BAR(2)$	0.967***	0.981***	0.919***	0.930***
$BVAR_W$	1.033	1.075	1.012	1.059
$BVAR_{W,R}$	0.814***	0.803***	0.846***	0.897***

Notes:
- The table reports (i) the value of the RMSEs for the benchmark model BAR(1) and (ii) the ratio between the RMSE of each of the other models and the RMSE of the AR benchmark.
- ***, ** and * indicate that the RMSE ratios are significantly different from 1 at the significance levels of 1%, 5% and 10%, respectively, according to the Diebold-Mariano t-statistic test for equal RMSEs.

Table 3: Point forecasts for Norway. Normal-Wishart prior

h	3	6	9	12
$BAR(1)$	22.161	22.880	20.904	21.455
$BAR(2)$	0.911*	0.898**	0.927*	0.891**
$BVAR_W$	1.001	1.019	1.036	1.022
$BVAR_{W,R}$	0.827**	0.851**	0.803***	0.842**

Notes:
- The table reports (i) the value of the RMSEs for the benchmark model BAR(1) and (ii) the ratio between the RMSE of each of the other models and the RMSE of the AR benchmark.
- ***, ** and * indicate that the RMSE ratios are significantly different from 1 at the significance levels of 1%, 5% and 10%, respectively, according to the Diebold-Mariano t-statistic test for equal RMSEs.

Table 4: Point forecasts for Canada. Minnesota prior

h	3	6	9	12
$BAR(1)$	12.401	13.762	13.890	13.069
$BAR(2)$	0.915***	0.905***	0.933***	0.924***
$BVAR_W$	1.012	1.041	1.070	1.033
$BVAR_{W,R}$	0.907***	0.881***	0.883***	0.915***

Notes:
- The table reports (i) the value of the RMSEs for the benchmark model BAR(1) and (ii) the ratio between the RMSE of each of the other models and the RMSE of the AR benchmark.
- ***, ** and * indicate that the RMSE ratios are significantly different from 1 at the significance levels of 1%, 5% and 10%, respectively, according to the Diebold-Mariano t-statistic test for equal RMSEs.

Table 5: Point forecasts for the UK. Minnesota prior

h	3	6	9	12
$BAR(1)$	7.911	8.207	7.944	8.103
$BAR(2)$	0.995**	0.988***	0.979***	0.981***
$BVAR_W$	1.015	1.004	1.009	1.046
$BVAR_{W,R}$	0.991***	0.989***	0.971***	0.970***

Notes:
- The table reports (i) the value of the RMSEs for the benchmark model BAR(1) and (ii) the ratio between the RMSE of each of the other models and the RMSE of the AR benchmark.
- ***, ** and * indicate that the RMSE ratios are significantly different from 1 at the significance levels of 1%, 5% and 10%, respectively, according to the Diebold-Mariano t-statistic test for equal RMSEs.

Table 6: Point forecasts for Norway. Minnesota prior

h	3	6	9	12
$BAR(1)$	16.002	16.738	16.150	16.594
$BAR(2)$	0.998**	0.972***	0.970***	0.983***
$BVAR_W$	1.037	1.030	1.018	1.029
$BVAR_{W,R}$	0.873***	0.905***	0.897***	0.886***

Notes:
- The table reports (i) the value of the RMSEs for the benchmark model BAR(1) and (ii) the ratio between the RMSE of each of the other models and the RMSE of the AR benchmark.
- ***, ** and * indicate that the RMSE ratios are significantly different from 1 at the significance levels of 1%, 5% and 10%, respectively, according to the Diebold-Mariano t-statistic test for equal RMSEs.

Table 7: Density forecasts for Canada. Normal-Wishart prior

h	3	6	9	12
$BAR(1)$	-8.709	-9.014	-8.943	-8.816
$BAR(2)$	0.032***	0.024***	0.005**	0.041***
$BVAR_W$	-0.007	-0.016	-0.049	-0.008
$BVAR_{W,R}$	0.044***	0.073***	0.097***	0.066***

Notes:
- The table reports (i) the value of the average LPSs for the benchmark model BAR(1) and (ii) the diffrence between the RMSE of each of the other models and the RMSE of the AR benchmark.
- ***, ** and * indicate that the RMSE ratios are significantly different from 1 at the significance levels of 1%, 5% and 10%, respectively, according to the Diebold-Mariano t-statistic test for equal RMSEs.

Table 8: Density forecasts for the UK. Normal-Wishart prior

h	3	6	9	12
$BAR(1)$	-6.264	-7.821	-6.907	-7.062
$BAR(2)$	0.063***	0.059***	0.038***	0.000
$BVAR_W$	-0.028	-0.010	-0.024	-0.052
$BVAR_{W,R}$	0.121***	0.116***	0.135***	0.102***

Notes:
- The table reports (i) the value of the average LPSs for the benchmark model BAR(1) and (ii) the diffrence between the RMSE of each of the other models and the RMSE of the AR benchmark.
- ***, ** and * indicate that the RMSE ratios are significantly different from 1 at the significance levels of 1%, 5% and 10%, respectively, according to the Diebold-Mariano t-statistic test for equal RMSEs.

Table 9: Density forecasts for Norway. Normal-Wishart prior

h	3	6	9	12
$BAR(1)$	-4.472	-5.369	-5.018	-5.105
$BAR(2)$	0.006*	0.013**	0.009*	0.018**
$BVAR_W$	-0.001	-0.001	-0.004	0.000
$BVAR_{W,R}$	0.055***	0.039***	0.040***	0.026***

Notes:
- The table reports (i) the value of the average LPSs for the benchmark model BAR(1) and (ii) the diffrence between the RMSE of each of the other models and the RMSE of the AR benchmark.
- ***, ** and * indicate that the RMSE ratios are significantly different from 1 at the significance levels of 1%, 5% and 10%, respectively, according to the Diebold-Mariano t-statistic test for equal RMSEs.

Table 10: Density forecasts for Canada. Minnesota prior

h	3	6	9	12
$BAR(1)$	-3.960	-4.559	-4.374	-4.008
$BAR(2)$	0.000	0.012**	0.003*	0.002*
$BVAR_W$	-0.014	-0.009	0.000	-0.018
$BVAR_{W,R}$	0.006**	0.010**	0.011***	0.004**

Notes:
- The table reports (i) the value of the average LPSs for the benchmark model BAR(1) and (ii) the diffrence between the RMSE of each of the other models and the RMSE of the AR benchmark.
- ***, ** and * indicate that the RMSE ratios are significantly different from 1 at the significance levels of 1%, 5% and 10%, respectively, according to the Diebold-Mariano t-statistic test for equal RMSEs.

Table 11: Density forecasts for the UK. Minnesota prior

h	3	6	9	12
$BAR(1)$	-3.012	-3.909	-3.827	-3.819
$BAR(2)$	0.025***	0.013**	0.017**	0.016**
$BVAR_W$	-0.006	0.000	-0.022	0.010
$BVAR_{W,R}$	0.020***	0.024***	0.031***	0.049***

Notes:
- The table reports (i) the value of the average LPSs for the benchmark model BAR(1) and (ii) the diffrence between the RMSE of each of the other models and the RMSE of the AR benchmark.
- ***, ** and * indicate that the RMSE ratios are significantly different from 1 at the significance levels of 1%, 5% and 10%, respectively, according to the Diebold-Mariano t-statistic test for equal RMSEs.

Table 12: Density forecasts for Norway. Minnesota prior

h	3	6	9	12
$BAR(1)$	-3.007	-5.622	-3.237	-3.199
$BAR(2)$	-0.008	0.100***	0.073***	0.051***
$BVAR_W$	-0.009	-0.012	-0.137	0.082
$BVAR_{W,R}$	0.065***	0.000	-0.015	0.070***

Notes:
- The table reports (i) the value of the average LPSs for the benchmark model BAR(1) and (ii) the diffrence between the RMSE of each of the other models and the RMSE of the AR benchmark.
- ***, ** and * indicate that the RMSE ratios are significantly different from 1 at the significance levels of 1%, 5% and 10%, respectively, according to the Diebold-Mariano t-statistic test for equal RMSEs.

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Conclusions

This thesis empirically investigates the business cycle dynamics in small open economies. Employing the more relevant indicators at business cycle frequency for this type of economies, the main messages are essentially two: the dependence of small open economies to global developments and the commonalities which are shown to share. The two outcomes come out from all the three papers which composed this thesis, under two different empirical point of view; indeed in the first chapter the transmission mechanism of global shocks highlights the importance of international fluctuations for a commodity-exporting country (Australia) and the beneficial effects of resource price surge on national real economic variables. The second chapter shows that economic variables employed to predict incoming recessions behave in the same way when a slumps is forthcoming; moreover it is verified that global factors influence and actually increase the forecasting performance of the early warning system. Finally in the third chapter the dependence of inflation dynamics to global and regional price level is tested; on one hand it is found that the world development are not enough to produce more reliable evidence on national inflation rate when using data at monthly frequency; on the other hand, when measures of regional inflation are included the results testify the increasing forecasting performance with respect to the baseline specifications.