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Fiscal Policy, Spillovers and Inflation in the Euro Area

Politique budgétaire, retombées et inflation dans la zone euro

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Jury

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Introduction générale

Cette thèse a pour objet empirique la zone euro. Elle traite de trois défis macroéconomiques d'une importance capitale pour le fonctionnement de la zone euro. Ces défis découlent de la structure institutionnelle particulière de la zone euro. Cette configuration est décrite par une union monétaire avec une autorité monétaire unique dont l'objectif principal est d'assurer la stabilité des prix et plusieurs autorités fiscales qui mènent la politique budgétaire de manière décentralisée. Des règles communes guident la politique budgétaire afin d'empêcher les pays de la zone euro d'atteindre des niveaux d'endettement insoutenables, ce qui évite à l'autorité monétaire de devoir monétiser la dette (Sargant and Wallace, 1981; Woodford, 2001). Il n'existe pas de capacité de stabilisation budgétaire permanente au niveau de la zone euro. Kirsanova et al. (2009) appellent cette répartition des responsabilités macroéconomiques le *consensus assigné*. La politique monétaire se concentre sur le contrôle de l'inflation, tandis que la politique budgétaire se concentre sur le contrôle des niveaux de dette publique. Cependant, cette séparation des tâches ne signifie pas que les politiques monétaire et budgétaire ne s'influencent pas ; elles interagissent l'une avec l'autre par de multiples canaux (ECB, 2021). En bref, la politique monétaire influence directement les conditions financières et donc les coûts d'emprunt de l'État, et plus indirectement la demande agrégée et les prix, et donc les recettes et les dépenses de l'État. La politique budgétaire influence le niveau des prix à la fois par le biais de la demande agrégée et par son impact sur l'offre. Elle influence également les conditions financières par le biais de l'émission de dette. Outre le cadre monétaire et budgétaire, un autre principe directeur de la zone euro est le marché unique, qui garantit la libre circulation des biens, des services, des capitaux et des personnes dans l'Union européenne. Ces canaux ne permettent toutefois pas un ajustement macroéconomique complet en raison des différences structurelles et institutionnelles (et des obstacles) existant entre les pays de la zone euro. Le défi auquel est confrontée la zone euro est donc, comme le dit si bien Bartsch et al. (2020), "de trouver un cadre institutionnel efficace (et un équilibre politique) qui puisse simultanément assurer (i) une stabilisation efficace au niveau national, et (ii) une orientation monétaire appropriée au niveau de la zone euro"[1]. Les facteurs et chocs mondiaux, les fluctuations cycliques et les interdépendances de l'action politique constituent un défi pour trouver cet équilibre. Cette thèse s'intéresse à certains de ces défis.

[1] Traduction de la citation de l'anglais.

Le marché unique et la monnaie unique ont renforcé l'intégration commerciale et financière entre les pays de la zone euro (voir, par exemple, Berger and Nitsch, 2008 pour une analyse de l'effet commercial de l'intégration de l'euro). Les échanges (importations et exportations) au sein de la zone euro représentaient environ 43% du total des échanges et 48% du PIB en 2022. Cette intégration augmente des retombées de la politique macroéconomique d'un pays sur les performances économiques d'autres pays et de la zone euro dans son ensemble. J'analyse l'ampleur des retombées de la politique budgétaire dans la zone euro dans le deuxième chapitre. Blanchard et al. (2021) parlent dans ce contexte de *externalités de la demande*[2]. Les politiques budgétaires sont décidées et financées au niveau national dans la zone euro, tandis que les bénéfices associés sont partiellement partagés avec d'autres pays dans un marché intégré (voir également ECB, 2021). Plus l'ampleur de ces *externalités de la demande* augmente, plus le risque augmente que la politique budgétaire soit sous-utilisée en tant qu'outil de stabilisation au niveau national et donc au niveau de la zone euro dans son ensemble, car une partie croissante de l'effet de la demande sera exportée (Blanchard et al., 2021). Je montre au chapitre 2 que l'ampleur des retombées n'est pas négligeable dans la zone euro, mais qu'elle dépend fortement du degré d'intégration et de la taille des pays. Les retombées ne sont pas seulement importantes au niveau des pays, mais aussi pour l'évaluation de l'orientation budgétaire de la zone euro dans son ensemble, qui affecte la demande agrégée et le niveau agrégé des prix et, par conséquent, la conduite de la politique monétaire par la Banque centrale européenne (BCE). La prise en compte de ces retombées pour la politique monétaire et la politique budgétaire nationale est donc très importante pour la zone euro.

Les pays qui ne font pas partie d'une union monétaire disposent d'instruments monétaires et budgétaires, en plus de la flexibilité du taux de change, pour lisser le cycle et amortir les chocs macroéconomiques. La politique monétaire et la flexibilité du taux de change jouent un rôle crucial dans la réponse aux chocs communs (globaux) (Friedman, 1953; Mundell, 1961 et plus récemment Obstfeld, 2020). Toutefois, au sein d'une union monétaire où la politique monétaire commune se concentre sur la situation globale, la politique budgétaire doit jouer un rôle plus important dans l'absorption des chocs macroéconomiques idiosyncratiques (propres à chaque pays) et assurer la stabilisation macroéconomique au niveau national. Même au-delà, la politique budgétaire peut être importante parce qu'elle est plus efficace que la politique monétaire après certains types de chocs qui affectent certains ménages, entreprises ou secteurs plus que d'autres, étant donné qu'elle peut généralement être mieux ciblée (voir ECB, 2021; Bartsch et al., 2020). La politique budgétaire peut également être activée lorsque la politique monétaire a largement épuisé sa marge de manœuvre à la borne inférieure effective. Les multiplicateurs budgétaires ont tendance à être plus élevés à la borne inférieure effective (Bonam et al.,

[2] Blanchard et al. (2021) propose ce terme en référence aux *externalités de la dette* plus communément connues (effets négatifs de la dette souveraine insoutenable d'un pays membre sur les autres pays membres, soit par les retombées des crises budgétaires, soit par la domination budgétaire de la politique monétaire, voir Blanchard et al., 2021).

2022). La politique budgétaire joue donc un rôle essentiel dans la stabilisation macroéconomique, en général et dans la zone euro en particulier (Corsetti et al., 2019 fournit un examen récent de la fonction de stabilisation de la politique budgétaire). J'aborde ce sujet dans le premier chapitre. J'y étudie dans quelle mesure les pays de la zone euro, en moyenne et individuellement, ont mené des politiques budgétaires contracycliques pour amortir les fluctuations macroéconomiques depuis le début de l'Union économique et monétaire (UEM). Je fais la distinction entre l'intention et le résultat des politiques budgétaires au cours du cycle, ainsi qu'entre les périodes normales et les périodes de crise (plus précisément, la grande crise financière et la crise de la dette européenne). Cette distinction entre les intentions et les résultats permet d'étudier les raisons pour lesquelles la politique budgétaire ne parvient pas à fournir les efforts de stabilisation nécessaires dans certaines situations. La littérature empirique n'a pas encore été en mesure de répondre pleinement à la question des raisons d'une politique budgétaire procyclique jusqu'à présent.[3] La conduite de politiques budgétaires contracycliques est importante du point de vue de chaque pays, mais aussi du point de vue de la zone euro dans son ensemble, car les politiques budgétaires individuelles affectent l'orientation budgétaire globale et donc le niveau agrégée des prix et l'espace de politique monétaire.

Les données utilisées dans ce chapitre se terminent en 2019, avant la crise de Covid. Pourtant, le sujet reste d'actualité, avec la crise de Covid et la crise de l'énergie qui ont suivi, et les vastes programmes de soutien budgétaire mis en œuvre dans la zone euro et ailleurs dans le monde pour amortir l'impact de ces chocs (voir ECB, 2021 et Haroutunian et al., 2021 pour une vue d'ensemble des mesures initiales de soutien budgétaire dans la zone euro). Il est généralement admis que la combinaison d'une réponse budgétaire forte et d'une réaction importante de la politique monétaire a permis d'amortir l'impact de ces chocs sur l'économie de la zone euro.[4] Toutefois, plus récemment, il est également apparu qu'une politique budgétaire de soutien (procyclique)[5] peut compliquer la tâche de la BCE qui consiste à ramener l'inflation à son niveau cible.[6]

Enfin, le troisième chapitre adopte un point de vue plus agrégée et analyse les facteurs qui déterminent la dynamique de l'inflation dans la zone euro. Plus particulièrement, j'analyse l'influence de l'activité réelle domestique et des facteurs globaux sur l'inflation de la zone euro, en utilisant une approche de la courbe de Phillips. Ce chapitre se concentre à

[3] Dans une tentative plus récente d'analyser cette question, Gootjes and de Haan, 2023 constatent qu'un accès insuffisant au financement national ou international rend les politiques budgétaires plus procycliques, y compris pour les pays développés. J'ai mis en évidence une consolidation forcée due à la tension du marché et un effet de signalisation.

[4] Voir, par exemple, Schnabel, I. (2022), "Finding the right mix : monetary-fiscal interaction at times of high inflation", discours prononcé lors de la conférence des observateurs de la Banque d'Angleterre, le 24 novembre 2022.

[5] Ou un retrait insuffisant d'une politique fiscale accommodante.

[6] Citation en anglais: "Fiscal support measures to shield the economy from the impact of high energy prices should be temporary, targeted and tailored [...]. Fiscal measures falling short of these principles are likely to exacerbate inflationary pressures, which would necessitate a stronger monetary policy response.", BCE, Déclaration de politique monétaire, le 15 décembre 2022.

la fois sur la compréhension ex post des facteurs qui influent sur l'inflation et sur les prévisions. Comprendre les causes de l'inflation et prévoir sa dynamique est, bien sûr, d'une importance capitale pour l'autorité de politique monétaire afin de conduire sa politique monétaire. Cette question se pose pour la zone euro et la BCE, comme pour les autres banques centrales, quelle que soit la structure institutionnelle. La question revêt toutefois une importance particulière pour la zone euro, et ce pour deux raisons : Premièrement, la zone euro est globalement une économie très ouverte, ce qui accroît l'influence potentielle des facteurs mondiaux sur l'inflation domestique. Le commerce international hors zone euro représentait 63% du PIB de la zone euro en 2022. Deuxièmement, la zone euro est un importateur net d'énergie à grande échelle, ce qui l'expose fortement aux fluctuations des prix de l'énergie et d'autres matières premières sur les marchés mondiaux. Le ratio de dépendance énergétique de la zone euro, qui indique dans quelle mesure un pays ou une région dépend des importations d'énergie, s'élevait à 60% en 2021. Cela différencie par exemple la zone euro de la situation des États-Unis, dont l'ouverture aux échanges commerciaux n'est que de 25% du PIB (chiffre pour 2021) et dont la balance commerciale en matière d'énergie est globalement équilibrée. Il est donc important d'étudier l'influence des facteurs globaux en mettant l'accent plus particulièrement sur la zone euro. Dans ce chapitre, je cherche également à déterminer si le processus d'inflation dans la zone euro diffère selon les régimes d'inflation, c'est-à-dire s'il est plus rigide lorsque l'inflation est faible et plus dynamique lorsque l'inflation est élevée, en utilisant une courbe de Phillips par quantile.

Le troisième chapitre a été publié au début de 2018 et les données vont jusqu'en 2016. Si les données se concentrent donc sur l'épisode de faible inflation dans la zone euro dans les années 2010 ("missing inflation", voir Ciccarelli and Osbat, 2017), les conclusions restent pertinentes pour l'épisode actuel d'inflation élevée. Premièrement, la compréhension du rôle des facteurs externes et internes de l'inflation reste une question centrale pour la politique monétaire (voir, par exemple, Lane, 2022). Deuxièmement, les facteurs externes, en particulier les prix mondiaux des matières premières et de l'énergie, ont fortement contribué à la hausse initiale de l'inflation globale fin 2021 (Lane, 2022 ; Nickel et al., 2022). Les modèles standards, tels que la courbe de Phillips, avec des hypothèses standards, n'ont toutefois pas été en mesure de prévoir avec précision cette forte augmentation de l'inflation, compte tenu de la forte surprise liée à l'évolution des prix des des matières premières, ce qui a entraîné des erreurs de prévision élevées (voir, par exemple, Chahad et al., 2022, qui concluent que "les erreurs d'inflation énergétique ont joué un rôle important dans les récentes sous-estimations de l'inflation"[7] dans la zone euro). Troisièmement, l'inflation core a augmenté plus progressivement au cours de l'épisode inflationniste récent et les prévisions de court terme étaient plus précises. Quatrièmement, le processus d'inflation s'est avéré très dynamique, conformément à la conclusion empirique que j'ai tirée de la courbe de Phillips quantile, selon laquelle l'inflation réagit

[7] Traduction de la citation de l'anglais

davantage aux pressions de la demande domestique et, en partie, aux facteurs mondiaux durant les épisodes de forte inflation que durant les épisodes de faible inflation. Harding et al. (2023) ont depuis formalisé ce résultat dans un modèle structurel. Ils proposent un modèle avec une courbe de Phillips qui est plate lorsque les pressions inflationnistes sont faibles et qui s'accentue lorsque les pressions inflationnistes augmentent, déterminée par une courbe de demande quasi coudée. Les non-linéarités s'appliquent surtout aux chocs de coûts (offre) (par rapport aux chocs de technologie et de demande). Une implication politique importante de ce résultat est que la politique monétaire, tout en étant plus efficace, est confrontée à un arbitrage plus sévère entre la stabilisation de l'inflation et de l'activité lorsque l'inflation est élevée. Cinquièmement, la forte hausse de l'inflation, due aux prix mondiaux des matières premières, a également mis en évidence l'interaction avec la politique budgétaire, qui a mis en œuvre des mesures directes de contrôle des prix dans de nombreux pays de la zone euro, interagissant ainsi directement avec les prix agrégés et donc la politique monétaire[8].

Plus généralement, l'identification de la courbe de Phillips reste d'actualité dans la littérature économique. Des travaux empiriques récents visent à améliorer le fondement microéconomique de la courbe de Phillips agrégée, en utilisant les coûts marginaux comme variable d'activité réelle, conformément à la théorie sous-jacente (Gali and Gertler, 1999).[9] Gagliardone et al. (2023) utilisent des données individuelles au niveau de l'entreprise et du produit pour montrer que la pente de la courbe de Phillips basée sur le coût marginal est élevée, ce qui suggère une répercussion importante des coûts marginaux sur l'inflation, mais que le lien entre les coûts marginaux et l'écart de production est relativement faible. La courbe de Phillips basée sur le coût marginal peut également constituer un meilleur outil que la courbe de Phillips basée sur l'écart de production pour retracer l'impact des chocs structurels, tels que les chocs d'offre, sur l'inflation. D'autres travaux, tels que ceux de Hooper et al. (2020) et Hazell et al. (2022), estiment les courbes de Phillips régionales, ce qui permet de niveler les effets du régime monétaire à long terme qui pourraient autrement influencer la pente de la courbe de Phillips. Le diagnostic à travers la courbe de Phillips reste bien "vivant"[10].

Cette thèse traite donc de plusieurs facteurs, tels que les retombées, les politiques nationales contracycliques et les chocs mondiaux, qui influencent la recherche d'un équilibre macroéconomique et la stabilisation macroéconomique simultanément au niveau de

[8] Dao et al. (2023) appellent ces mesures "politique budgétaire non conventionnelle", qui sont motivées par la volonté d'atténuer directement les effets de la hausse des prix de l'énergie sur l'inflation.

[9] Gagliardone et al. (2023) montrent que la formulation conventionnelle de la courbe de Phillips avec l'écart de production comme variable d'activité réelle ne tient que s'il existe une relation proportionnelle entre les coûts marginaux et l'écart de production. Même si cette condition est remplie, la pente de l'écart de production est en fin de compte le produit de deux paramètres : l'élasticité de l'inflation par rapport aux coûts marginaux réels et l'élasticité des coûts marginaux par rapport à l'écart de production (voir Gagliardone et al., 2023).

[10] Cette citation fait référence à une discussion de longue date dans la littérature sur la courbe de Phillips (citations en anglais) : "The Phillips curve is alive and well" (Gordon, 2013) et "The Phillips curve is dead".

la zone euro et au niveau national.

La description suivante donne un résumé plus approfondi de chacun des chapitres.

Dans le premier chapitre, j'étudie l'orientation budgétaire sous la forme du "stance" dans la zone euro en temps réel pour la période allant de 1999 à 2019. L'orientation budgétaire reflète la réaction discrétionnaire, c'est-à-dire non automatique, du solde budgétaire aux conditions cycliques. L'orientation budgétaire est considérée comme contracyclique si la partie non automatique du solde budgétaire va dans le même sens que l'activité (par exemple, en détériorant le solde pour faire face aux ralentissements économiques), procyclique si les deux évoluent dans des directions opposées et a-cyclique si la partie non automatique du solde budgétaire ne réagit pas aux conditions cycliques. J'examine cette question séparément pour la phase de planification du budget, montrant les intentions de la politique budgétaire, et pour la phase de mise en œuvre, en utilisant une nouvelle base de données en temps réel. Je constate qu'en moyenne, pour les 11 pays de la zone euro considérés, l'orientation budgétaire est procyclique pendant la planification du budget et a-cyclique dans les résultats budgétaires. La tendance à mener une politique procyclique est donc déjà ancrée dans les plans budgétaires et n'est pas le résultat de surprises sur les conditions cycliques. J'utilise à la fois un estimateur de panel à effet fixe et une approche de variable instrumentale comme stratégie empirique. Les résultats sont robustes aux différentes spécifications et méthodes d'estimation. Ils sont également robustes lorsqu'on tient compte des deux périodes de crise, la crise financière mondiale et la crise de la dette européenne, qui ont frappé la zone euro au cours de la période couverte. Les plans budgétaires étaient également procycliques en dehors des années de crise, et contracycliques (mais non significatifs) pendant la grande crise financière et significativement procycliques pendant la crise de la dette européenne. Une explication possible de la procyclicité des plans budgétaires pourrait être que les gouvernements signalent une orientation politique plus stricte que celle qui est finalement mise en œuvre aux marchés. Je montre que la procyclicité lors de la phase de planification budgétaire est plus fréquente et plus importante lors des épisodes de resserrement, alors que l'orientation budgétaire tend à être plus a-cyclique lors des épisodes d'assouplissement budgétaire en moyenne. Je détecte également une forte hétérogénéité des pays dans l'orientation de la politique budgétaire dans la zone euro. J'observe un comportement plus procyclique des plans budgétaires pour des pays tels que la Grèce, la France, l'Italie, l'Espagne et la Finlande et un comportement plus contracyclique des plans budgétaires pour un groupe de pays comprenant l'Irlande, l'Allemagne, l'Autriche et, dans une moindre mesure, les Pays-Bas et la Belgique. L'hétérogénéité de la cyclicité de la politique budgétaire est importante au sein de la zone euro.

Dans le deuxième chapitre, je présente de nouvelles estimations empiriques de l'ampleur des retombées budgétaires dans la zone euro pour la période allant de 1972 à 2017. Les retombées sont définies comme une partie de l'action budgétaire dans un ensemble des pays qui affecte l'activité, les prix et d'autres variables macroéconomiques dans d'autres pays. Dans le contexte de l'UEM, les retombées budgétaires sont particulièrement importantes pour l'évaluation précise des perspectives économiques dans la zone euro, ainsi que pour le débat sur un changement coordonné de l'orientation budgétaire et sur une capacité budgétaire commune de la zone euro. Le chapitre se concentre sur les retombées de l'action budgétaire basée sur les dépenses dans la perspective du pays de destination. Il montre l'ampleur des retombées des mesures budgétaires prises dans tous les pays, sauf un, sur le pays restant. J'utilise un modèle vectoriel autorégressif de panel qui complète les variables macroéconomiques nationales pertinentes par des dépenses publiques étrangères pondérées en fonction des échanges commerciaux. Je constate que les dépenses budgétaires ont des retombées positives et non négligeables entre les pays de la zone euro. L'ampleur des retombées budgétaires est toutefois inférieure à celle des multiplicateurs budgétaires domestiques. L'effet est en moyenne d'environ 0,4 après deux ans, par rapport à un multiplicateur fiscal domestique moyen pour les dépenses publiques d'environ 0,9. L'ampleur des retombées dépend essentiellement de la taille et de l'ouverture des pays d'où provient le choc fiscal. Les économies plus petites et plus ouvertes génèrent des retombées plus importantes que les économies plus grandes et plus fermées. Je montre également que la taille et la persistance des retombées fiscales ont légèrement augmenté au cours de la période analysée.

Le dernier chapitre étudie le rôle des facteurs mondiaux dans l'inflation de la zone euro pour la période 1996-2016 à l'aide de la courbe de Phillips. Avec l'intégration économique croissante, les facteurs mondiaux peuvent jouer un rôle de plus en plus important dans l'inflation domestique. Il est important de comprendre ces facteurs afin d'analyser et de prévoir avec précision la dynamique de l'inflation domestique. J'examine un large éventail d'indicateurs mondiaux, des indicateurs mondiaux plus traditionnels tels que les prix des matières premières, les taux de change et les prix à l'importation, jusqu'aux indicateurs proposés plus récemment dans la littérature, tels que les prix mondiaux à la consommation et l'écart de production mondial. Je constate que les indicateurs traditionnels des prix des matières premières et des prix à l'importation permettent une bonne identification de la courbe de Phillips pour l'inflation de la zone euro, contrairement aux mesures de l'écart de production mondial proposées par Auer et al. (2017), qui n'ont pas d'influence significative sur l'inflation de la zone euro. Les indicateurs mondiaux n'ont pas non plus de bonnes propriétés pour la prévision de l'inflation en zone euro. Je montre plus généralement que les performances prévisionnelles de la courbe de Phillips dépendent fortement de la période considérée. Je montre également que la courbe de Phillips pourrait ne pas être linéaire en analysant l'ensemble de la distribution conditionnelle de l'inflation, et pas seulement la moyenne, à l'aide d'une approche de régression quantile dynamique. Je constate que le processus d'inflation est plus persistant dans la queue gauche de la distribution, c'est-à-dire lorsque l'inflation se situe dans ses quantiles inférieurs. En revanche, l'activité domestique - et, dans une moindre mesure, les facteurs mondiaux - ont une influence plus forte sur l'inflation pendant les périodes d'inflation plus élevée. Par conséquent, les

régressions par quantile peuvent améliorer le pouvoir prédictif de la courbe de Phillips pendant certaines périodes de faible inflation persistante (2014-2015).

Mots-clés: Politique budgétaire, multiplicateurs budgétaires, retombées budgétaires, dépenses publiques, stabilisation macroéconomique, procyclicité, inflation, zone euro, prévision, données en temps réel, autorégressions vectorielles, courbe de Phillips, régressions quantile.

JEL classifications: C22, C33, C36, E31, E37, E62, E63, F45, H50

General Introduction

This thesis has the euro area as empirical focus. It deals with three macroeconomic challenges of central importance to the functioning of the euro area. These challenges arise from the special institutional setup of the euro area. This setup is described by a monetary union with a single monetary authority whose primary objective is to deliver price stability and several fiscal authorities who conduct fiscal policy in a decentralised manner[11]. Common rules guide fiscal policy to prevent euro area countries from reaching unsustainable debt levels, which in turn guards the monetary authority from having to monetise debt (Sargant and Wallace, 1981; Woodford, 2001). There is no permanent fiscal stabilisation capacity at the euro area level. Kirsanova et al. (2009) call this division of macroeconomic responsibilities the *consensus assignment*. Monetary policy focuses on inflation control, while fiscal policy focuses on the control of government debt levels. However, this separation of tasks does not mean that monetary and fiscal policy do not influence each other; they interact with each other through multiple channels (ECB, 2021). In brief, monetary policy influences financial conditions and hence government borrowing costs directly, and aggregate demand and prices and hence government revenues and expenditures more indirectly. Fiscal policy influences the price level both through aggregate demand and through its impact on the supply side. It also influences financial conditions through debt issuance. Apart from this monetary and fiscal framework, a further guiding principle of the euro area is the single market, which ensures free movement of goods, services, capital and persons in the European Union. These channels do not allow, however, for a full macroeconomic adjustment due to remaining structural and institutional differences (and barriers) between euro area countries. A main challenge faced by the euro area is hence, as Bartsch et al. (2020) appropriately put it, "to find an effective institutional framework (and political equilibrium) that can simultaneously ensure (i) efficient stabilisation at national level, and (ii) the appropriate monetary stance at the euro area level". Global factors and global shocks, cyclical swings and interdependencies of political action constitute a challenge for finding this equilibrium. This thesis deals with some of them.

The single market and the single currency have reinforced trade and financial integration between euro area countries (see, for example, Berger and Nitsch, 2008 for an

[11] The principles of this setup are laid down in the Treaty on the Functioning of the European Union (TFEU).

analysis of the trade effect of euro integration). Trade (imports and exports) within the euro area accounted for about 43% of total trade and for about 48% of GDP in 2022. This integration creates the case for spillovers of one country's macroeconomic policy on the economic performance of other countries and the euro area as a whole. I analyse the magnitude of spillovers from fiscal policy in the euro area in chapter two. Blanchard et al. (2021) speak in this context of *demand externalities*[12]. Fiscal policies are decided and funded at the country level in the euro area, while the associated benefits are partly shared with other countries in integrated markets (see also ECB, 2021). As the size of these *demand externalities* increases, so does the risk that fiscal policy will be underused as a stabilisation tool at the country level and hence at the euro area level as a whole, as an increasing part of the demand effect will be exported (Blanchard et al., 2021). I show in chapter two that the size of spillovers is non-negligible in the euro area, but it depends greatly on the degree of integration and the size of countries. Spillovers are not only relevant at the country level but also for the assessment of the fiscal stance for the euro area as a whole, which affects aggregate demand and the aggregate price level and hence the conduct of monetary policy by the European Central Bank (ECB). The consideration of these spillovers for monetary policy and domestic fiscal policy is hence highly relevant in the euro area.

Countries that are not part of a monetary union have both monetary and fiscal instruments, in addition to exchange rate flexibility, to smooth the cycle and cushion macroeconomic shocks. Monetary policy and exchange rate flexibility play a crucial role in responding to common (aggregate) shocks (Friedman, 1953; Mundell, 1961 and more recently Obstfeld, 2020). However, within a currency union where the common monetary policy focuses on the aggregate situation, fiscal policy needs to play a greater role in absorbing idiosyncratic (country-specific) macroeconomic shocks and provide macroeconomic stabilisation at the country-level. And even beyond that, fiscal policy can be important because it is more effective than monetary policy after certain type of adverse shocks that affect some households, firms or sectors more than others, as it is found to be generally better suited to target policies (see ECB, 2021; Bartsch et al., 2020). Fiscal policy can also be activated when monetary policy has largely exhausted its policy room at the effective lower bound, given that fiscal multipliers tend to be higher at the effective lower bound (Bonam et al., 2022). Fiscal policy thus plays an essential role in macroeconomic stabilisation, in general and in the euro area in particular (Corsetti et al., 2019 provide a recent review of the stabilisation function of fiscal policy). I address this topic in chapter one. There, I study the extent to which euro area countries, on average and individually, have conducted counter-cyclical fiscal policies to cushion macroeconomic fluctuations since the start of the Economic and Monetary Union (EMU). I distinguish

[12] Blanchard et al. (2021) have coined this term in reference to the more commonly known debt externalities (adverse effects of unsustainable sovereign debt in one member country on other member countries, either through the spillovers of fiscal crises or through fiscal dominance of monetary policy, see Blanchard et al., 2021).

between the intention and the actual outcome of fiscal policies over the cycle, as well as between normal times and crisis periods (more precisely, the Great Financial crisis and the European Debt crisis). This distinction between intentions and outcomes allows to study the reasons why fiscal policy fails to deliver the necessary stabilisation efforts in some situations. The empirical literature has not been able to fully answer the question about the reasons for a pro-cyclical fiscal policy yet.[13] The conduct of counter-cyclical fiscal policies is important from an individual country perspective, but also from an aggregate euro area point of view, as individual fiscal policies affect the aggregate fiscal stance and hence the aggregate price level and the monetary policy space.

The data used in this chapter ends in 2019, before the COVID-19 crisis. Yet, the topic remains timely, with the COVID-19 and the Energy crisis that followed, and the large fiscal support packages implemented in the euro area and elsewhere in the world to cushion the impact of these sizeable shocks (see ECB, 2021 and Haroutunian et al., 2021 for an overview of initial fiscal support measures in the euro area). There is a basic agreement that the combination of the strong fiscal response and the forceful monetary policy reaction were successful in cushioning the impact of these shocks on the euro area economy.[14] However, more recently it became also clear that supportive (pro-cyclical) fiscal policy[15] can complicate the task of the ECB to bring inflation back to target.[16]

Finally, chapter three takes a more aggregate view and analyses the factors that shape euro area wide inflation dynamics. More in particular, I analyse the influence of domestic real activity and global factors on euro area inflation, using a Philips curve approach. The focus of this chapter lies both on the ex post assessment of inflation drivers and on forecasting. Understanding the causes of inflation and forecasting its dynamics is, of course, of paramount importance for the monetary policy authority in order to conduct its monetary policy. This question arises for the euro area and the ECB, as for other central banks, regardless of the institutional setup of the euro area. The question is however of particular relevance for the euro area for two reasons: First, the euro area is on aggregate a very open economy, increasing the potential influence of global factors on domestic inflation. Extra euro area trade made up 63% of GDP in the euro area in 2022. Second, the euro area is a large-scale net energy importer, exposing it strongly to energy and other commodity price fluctuations on world markets. The euro area's energy dependency ratio[17] stood at 60% in 2021. This differentiates the euro area for example

[13] In a more recent attempt to address this question, Gootjes and de Haan, 2023 find that poor access to domestic or international finance makes fiscal policies more pro-cyclical, also for developed countries. I put into evidence forced consolidation due to market stress and a signalling effect.

[14] See, for example, Schnabel, I. (2022), "Finding the right mix: monetary-fiscal interaction at times of high inflation", speech at the Bank of England Watchers' Conference, 24 November 2022.

[15] Or an insufficient withdrawal of accommodative fiscal policy.

[16] "Fiscal support measures to shield the economy from the impact of high energy prices should be temporary, targeted and tailored [...]. Fiscal measures falling short of these principles are likely to exacerbate inflationary pressures, which would necessitate a stronger monetary policy response.", ECB, Monetary Policy Statement, 15 December 2022.

[17] Ratio between net imports and gross available energy, which indicates the extent to which a country

from the U.S. situation, who has an openness to trade of only 25% of GDP (figure for 2021) and a broadly balanced energy trade balance. It is therefore important to study the influence of global factors with a special focus on the euro area. In this chapter, I also look at whether the inflation process in the euro area differs in different inflation regimes, i.e. whether it is stickier when inflation is low and more dynamic in a high inflation regime, using a quantile Phillips curve.

Chapter three has been published in the beginning of 2018 and the data goes up to 2016. While the data focuses hence on the low inflation episode in the euro area in the 2010s ("missing inflation", see Ciccarelli and Osbat, 2017), the conclusions remain relevant for the current episode of high inflation. First, understanding the role of external and domestic drivers of inflation continues to be a central question for monetary policy (see, for example, Lane, 2022). Second, external factors, especially global commodity and energy prices, contributed strongly to the initial rise in headline inflation in late 2021 (see Lane, 2022; Nickel et al., 2022 for the euro area). Standard models, such as the Phillips curve, with standard assumptions were however not able to accurately predict this strong increase in inflation, given the strong surprise on commodity price developments, leading to high forecast errors (see Chahad et al., 2022, who conclude that "energy inflation errors played an important role in recent inflation underestimations" in the euro area). Third, core inflation increased more gradually during the recent inflation episode and short-term forecasts had a higher accuracy. Fourth, the inflation process proved very dynamic, corroborating my empirical finding from the quantile Phillips curve that inflation responds stronger to domestic demand pressures and partly also to global factors during high inflation episodes compared to low inflation episodes. Harding et al. (2023) have since formalised this finding in a structural model. They propose a model with a Phillips curve that is flat when inflationary pressures are subdued and steepens as inflationary pressures rise, determined by a quasi-kinked demand schedule. The non-linearities apply especially to cost-push (supply) shocks (compared to technology and demand shocks). An important policy implication of this finding is that monetary policy, while being more effective, faces also a more severe trade-off between inflation and output stabilisation when inflation is high (Harding et al., 2023). Fifth, the sharp rise in inflation, driven by global commodity prices, has also exposed the interaction with fiscal policy, which has implemented direct price control measures in some euro area countries, interacting directly with aggregate prices and hence monetary policy[18].

More generally, the identification of the Phillips curve remains topical in the economic literature. Recent empirical work aims at a better micro-foundation of the aggregate Phillips curve, using marginal costs as real activity variable, in line with underlying theory (Gali and Gertler, 1999).[19] Gagliardone et al. (2023) use individual firm-product level

or a region is dependent on energy imports.

[18] Dao et al. (2023) call these measures "unconventional fiscal policy", which are motivated by a desire to directly mute the effects of the increase in energy prices on inflation.

[19] Gagliardone et al. (2023) show that the conventional formulation of the Phillips curve with the output

data to show that the slope of the marginal cost-based Phillips curve is high, suggesting a substantial pass-through of marginal costs into inflation, but that the connection between marginal costs and the output gap is relatively weak. The marginal cost-based Phillips curve can also provide a better tool than the output gap-based Phillips curve to trace the impact of structural shocks, such as supply shocks, on inflation. Other work, such as Hooper et al. (2020) and Hazell et al. (2022), estimate regional Phillips curves, which allow levelling out the effects of the long-run monetary regime that might otherwise influence the slope of the Phillips curve. The diagnosis across the Phillips curve remains well "alive"[20].

This thesis deals hence with several factors, such as spillovers, countercyclical national policies and global shocks, influencing on finding a macroeconomic equilibrium and macroeconomic stabilisation simultaneously at the euro area level and the national level.

The following overview gives a deeper summary of each of the chapters.

In chapter one, I study the fiscal orientation in form of the fiscal stance in the euro area in real-time for the period from 1999 to 2019. The fiscal stance reflects the discretionary, i.e. non-automatic, reaction of the budget balance to cyclical conditions. The fiscal stance is considered to be counter-cyclical if the non-automatic part of the budget balance goes in the same direction as the output (e.g. deteriorating the balance to face economic downturns), pro-cyclical if both evolve in opposite directions and a-cyclical if the non-automatic part of the budget balance does not respond to cyclical conditions. I examine this question separately for the budget planning phase, showing the intentions of fiscal policy makers, and the implementation phase, using a new real-time dateset. I find that, on average for the 11 euro area countries under consideration, the fiscal stance is pro-cyclical during budget planning and a-cyclical in fiscal outcomes. The tendency to run a pro-cyclical policy is hence already anchored in fiscal plans and not just an outcome of surprises on cyclical conditions. I use both a panel fixed effect estimator and an instrumental variable approach as empirical strategy. The results are robust to different specifications and estimation methods. They are also robust when controlling for the two crisis periods, the Global Financial Crisis and the European Debt Crisis, that hit the euro area during the covered period. Fiscal plans were also pro-cyclical outside of the crisis years, and counter-cyclical (but not significant) during the Great Financial Crisis and significantly pro-cyclical during the European Debt Crisis. A possible explanation for the pro-cyclicality of fiscal plans could be for governments to signal a tighter policy stance than is implemented in the end. I show that pro-cyclicality at the budget planning phase is more frequent and more significant during tightening episodes, whereas the fiscal stance tends to be more a-cyclical during fiscal loosening episodes on average. I also detect strong country heterogeneity in the orientation of fiscal policy in the euro area. I

gap as real activity variable only holds if there is a proportional relationship between marginal costs and the output gap. Even if this condition holds, the output-gap slope is ultimately the product of two parameters: the elasticity of inflation with respect to real marginal costs and the elasticity of marginal costs with respect to the output gap (see Gagliardone et al., 2023).

[20]This quotation refers to a long-standing discussion in the literature on the Phillips curve: "The Phillips curve is alive and well" (Gordon, 2013) and "The Phillips curve is dead".

observe a more pro-cyclical behaviour of fiscal plans for countries such as Greece, France, Italy, Spain, and Finland and a more counter-cyclical behaviour in fiscal plans for a group of countries covering Ireland, Germany, Austria, and to a lesser extent the Netherlands and Belgium. The heterogeneity of the cyclicality of fiscal policy is important in the euro area.

In the second chapter, I provide new empirical estimates of the magnitude of fiscal spillovers in the euro area for the period from 1972 to 2017. Spillovers are defined as part of the fiscal action in a given set of countries that affects output, prices and other macroeconomic variables in other countries. In the context of EMU, fiscal spillovers are particularly relevant for the accurate assessment of the cyclical situation in the euro area, as well as in the debate on a coordinated change in the euro area fiscal stance and on a common euro area fiscal capacity. The chapter focuses on spillovers from expenditure-based fiscal action in a destination country perspective. This shows the magnitude of spillovers from fiscal action in all but one country to the remaining country. I employ a panel vector autoregressive model that augments the relevant domestic macroeconomic variables with trade-weighted foreign government spending. I find that fiscal spending has positive and non-negligible spillovers across countries in the euro area. The magnitude of fiscal spillover lies however below those of domestic fiscal multipliers. The spillover effect is on average about 0.4 after two years, compared to an average domestic fiscal multiplier for government spending of about 0.9. The size of spillovers depends crucially on the size and the openess of the countries where the fiscal shock originates. Smaller and more open economies generate larger spillovers compared to larger and more closed economies. I also show that the size and persistency of fiscal spillovers have slightly increased over the period under consideration.

The last chapter studies the role of global factors in euro area inflation for the period from 1996 to 2016 with the help of the Phillips curve. With increasing economic integration, global factors may play an increasing role in domestic inflation. It is important to understand these factors in order to accurately analyse and predict domestic inflation dynamics. I examine a wide range of global indicators starting from more traditional global indicators such as commodity prices, exchange rates and import prices to indicators discussed more recently in the literature, such as global consumer inflation and global economic slack. I find that traditional commodity price and import price indicators provide a good identification of the Phillips curve for euro area inflation, in contrast to global economic slack measures proposed by Auer et al. (2017), which do not have a significant influence on euro area inflation. Global indicators also do not have good properties for forecasting domestic inflation. I show more generally that the forecast performance of the Phillips curve depends heavily on the considered time period. I also show that the Phillips curve might not be linear by analysing the entire conditional distribution of inflation and not just the mean using a dynamic quantile regression approach. I find that the inflation process is more persistent at the left tail of the distribution, i.e. when inflation is in its

lower quantiles. In contrast, domestic activity, and to a lesser extent also global factors, are found to have a stronger influence on inflation during periods of higher inflation. Consequently, quantile regressions can improve the forecast ability of the Phillips curve during some periods of persistently low inflation (2014-2015).

Keywords: Fiscal policy, fiscal multipliers, fiscal spillovers, government spending, macroeconomic stabilisation, procyclicality, inflation, euo area, forecasting, real-time data, vector autoregressions, Phillips curve, quantile regressions.

JEL classifications: C22, C33, C36, E31, E37, E62, E63, F45, H50

Publication Details

This thesis consists of three chapters. The information below gives the publication details and co-authorships for each of the chapters.

	Title	Author	Publication
Ch.1	Fiscal Policy Orientation in the Euro Area in Real-time	Katja Schmidt & Antoine Sigwalt (Banque de France, PSE)	Published as Working Paper (2022), Working Paper Series N°896, Banque de France
Ch.2	Fiscal Expenditure Spillovers in the Euro Area	Katja Schmidt (single author)	Published as part of Working Paper (2020), ECB Occasional Paper N°240, European Central Bank (full paper with M. Alloza, M. Ferdinandusse and P. Jacquinot)
Ch.3	Explaining and Forecasting Euro Area Inflation: the Role of Domestic and Global Factors	Sophie Bereau, (U. de Namur, U. de Lorraine) Violaine Faubert (Banque de France) & Katja Schmidt	Published as Working Paper (2018), Working Paper Series N°663, Banque de France

Chapter 1

Fiscal Policy Orientation in the Euro Area in Real-time[1]

1.1 Introduction

This chapter analyses the fiscal stance in euro area countries in real-time. We measure the fiscal stance as the change in the cyclically-adjusted (CAB) or the primary cyclically-adjusted (CAPB) budget balance. This measure reflects the discretionary, i.e. non-automatic, reaction of the budget balance to cyclical conditions. In our analysis, we compare the anticipated fiscal stance at the time of budget planning, which normally takes place in autumn of a given year for the following year, with its realisation a year after. We examine the orientation of fiscal policy throughout the cycle, i.e. its counter- or pro-cyclicality, both at the time of budget planning and at the time of the first outcome. We classify fiscal policy as counter-cyclical if the change in the cyclically-adjusted (primary) budget balance and the output gap have the same sign (i.e. a positive output gap goes along with a tightening of fiscal policy and a negative output gap with a loosening of fiscal policy), pro-cyclical if they have the opposite sign, and acyclical if they are broadly neutral.

One of the objectives of fiscal policy is to stabilise the economy along its potential (Musgrave, 1959). When economic activity falls short of its potential, policymakers should implement an expansionary fiscal stance to support demand and vice versa, beyond the automatic impact of cyclical conditions on the budget balance via automatic stabilisers. However, during the budget planning process, decision-makers are confronted with uncertainties about the exact cyclical position of the economy. Output gap estimates, which are frequently used to quantify the cycle, are surrounded by high estimation uncertainties, particularly in real-time. Indeed, they are based on the real-time values of potential GDP, an unobservable variable, and suffer from frequent and sometimes large ex-post revisions.

[1] This chapter, co-authored with A. Sigwalt (Banque de France and Paris School of Economics), has been published as Working Paper Banque de France in December, 2022.

The same uncertainties and ex-post revisions then also affect the cyclically-adjusted budget balance, which is derived from output gap measures. By construction, a revision of the output gap of a given sign (say, an increase), brings about a simultaneous revision of the cyclically-adjusted balance of the opposite sign (in this case, a deterioration), all other things being equal, given that the cyclically-adjusted balance is negatively linked to the output gap. Some authors suggest (see, for an overview, Cimadomo, 2016) that these output gap revisions may increase the degree of pro-cyclicality in ex-post data. This chapter aims to examine this question in distinguishing between the ex-ante and the ex-post assessment of the fiscal stance.

Several papers (Bankowski and Ferdinandusse, 2017; Commission, 2004; Fatas and Mihov, 2010) have analysed the question of the cyclicality of the fiscal stance in the euro area. They generally find a pro-cyclical fiscal orientation, at least periodically, of fiscal policy since the start of the European Monetary Union (EMU). Their assessment is based on the analysis of ex-post data, as they do not take into account the information actually available to policy-makers in real-time. Forni and Momigliano (2005), Golinelli and Momigliano (2009) and Cimadomo (2012) show however that estimating the fiscal stance based on ex-post revised data might give a misleading picture of the sensitivity of fiscal policies to cyclical conditions due to revisions of output gap estimates.

Uncertainty about the cyclical position is of course only one possible explanation for a pro-cyclical orientation of fiscal policy. The literature has proposed some other explanations: (1) a general deficit bias of fiscal policy due to political agency problems (see, among others, Alesina and Tabellini, 2005), which might lead to an insufficiently counter-cyclical fiscal policy in particular at the height of the cycle, (2) forced consolidation during periods of financial stress, where markets push countries to consolidate and thus encourage pro-cyclical behaviour during these periods, and (3) borrowing constraints which prevent countries to achieve sufficient counter-cyclical fiscal policy especially during bad macroeconomic times (see, among others, Gavin and Perotti, 1997). The later argument is especially present in the literature on developing countries and should play a smaller role in our panel - apart from isolated periods during the European Debt crisis. From a theoretical point of view, procyclical fiscal policy remains a puzzle.[2]

This chapter looks anew into the question of the cyclicality of fiscal policy for euro area countries in real-time over a long time horizon since the start of EMU in 1999 up to 2019, based on a newly constructed dataset from European Commission data. This time span includes notably the Great Financial crisis (GFC) and the European Debt crisis (EDC), during which the role of fiscal stabilisation have received renewed attention. The pro-cyclical nature of fiscal tightening during the European Debt crisis has been widely associated with an aggravation of the crisis. We examine whether fiscal policy turned out to be particularly pro-cyclical in real-time during the EDC.

[2]In neoclassical models, the optimal fiscal policy stance is either a-cyclical or counter-cyclical. In Keynesian models, on the other hand, with the presence of sticky prices, the optimal fiscal policy stance is counter-cyclical.

The rest of the chapter is organised as follows. Section 1.2 presents an overview of the relevant literature and how our research relates to it. Section 1.3 describes the data and the construction of the real-time dataset. Section 1.4 presents a descriptive analysis of the relationship between the fiscal stance and the cycle, while section 1.5 provides the empirical estimates. Finally, Section 1.6 concludes the paper.

1.2 Literature review

The literature on the assessment of the stance of fiscal policy over the business cycle has produced a large body of research, with sometimes contradicting results. We focus here on the literature for the euro area.

Ex-post analysis of the cyclicality tend to provide evidence of a pro-cyclical fiscal stance for euro area countries (Bankowski and Ferdinandusse, 2017; Commission, 2004). Since the work of Orphanides (2001), who found that substantial errors in the real-time assessment of cyclical conditions induce misleading conclusions on the stance of monetary policy, real-time assessment of cyclicality has also found its way into fiscal policy analysis. Forni and Momigliano (2005), Golinelli and Momigliano (2006) and Golinelli and Momigliano (2009) find a weakly counter-cyclical behaviour of the primary cyclically-adjusted balance with real-time data and a broadly a-cyclical behaviour with ex-post data for OECD countries. Cimadomo (2012) also shows a slightly counter-cyclical behaviour with real-time data, especially during buoyant economic times, that turns pro-cyclical ex-post. Beetsma and Giuliodori (2010) reach similar conclusions, showing that countries follow a more counter-cyclical fiscal policy at budget planning than during the implementation process. All four papers base their analysis on the December issues of the OECD Economic Outlook for the period between 1994 and the early 2000s and confirm hence that a misjudgment on the cyclical position weighs on the effectiveness of fiscal stabilisation. Pina (2009) examines the orientation of fiscal policy in real-time for euro area countries based on a database from the European Commission's autumn forecast. He confirms earlier results of a counter-cyclical orientation of planned fiscal policy, which turns a-cyclical ex-post, given that fiscal policy does not react significantly to updates in cyclical conditions. This contrasts with the results from Marinheiros (2008) based also on European Commission's data for the period from 1999 to 2006 who finds that discretionary fiscal policy has been designed to be pro-cyclical. Holm-Hadulla et al. (2010) undertake a similar analysis for one part of the budget balance, namely government expenditures, and find a pro-cyclical slippage in overall expenditure to surprises in the output gap. Cimadomo (2016) provides the most recent survey of the literature on the real-time assessment of fiscal policy.

In the most relevant work since the GFC, and using a large country panel of advanced and emerging economies, Poghosyan and Tosun (2019) find that discretionary fiscal policy in advanced economies is counter-cyclical at the planning stage, which disappears largely at budget implementation. Budget implementation even turned pro-cyclical after the GFC. Gootjes and de Haan (2022) find, for a sample of 27 EU countries over the 2000-2015 period, an a-cyclical orientation of fiscal policy at the planning stage and a slightly pro-cyclical policy at the implementation phase, basing their analysis on the European Commission's spring forecast. This result is however partly driven by non-euro area EU countries, while euro area EU countries are shown to have a higher degree of counter-cyclicality at the planning stage. Aldama and Creel (2022), based on a long sample of 19 OECD countries between 1997-2018, find, in contrast with the rest of the literature, a pro-cyclical fiscal orientation already at the planning stage. They do not analyse the realization of the fiscal policy orientation.

For EU countries in particular, the Stability and Growth Pact (SGP) recommends that fiscal consolidation should be counter-cyclical during expansion phases in order to meet the Medium-Term objective. Some commentators, however, have argued that the deficit of 3% or less enshrined in the SGP may force governments to cut expenditures or increase taxes during downturns, resulting in a pro-cyclical fiscal policy (see for instance Eichengreen and Wyplosz, 1998). This hypothesis seems to hold true using ex-post data: studies tend to show that fiscal policy outcomes have been at best a-cyclical or even pro-cyclical after the introduction of the euro (Cimadomo, 2012).

Conclusions regarding the cyclicality of fiscal policy are therefore not clear-cut in the literature and hinge on a variety of factors. On the one hand, the distinction between ex-ante plans and ex-post outcome is important. On the other hand, the definition of fiscal policy itself matters. The choice of using cyclically-adjusted balances or nominal balances is not benign (Bernoth et al., 2008). The former indeed requires an assessment of the output gap, which is unobservable and subject to revisions. In contrast, the nominal balance makes it difficult to assess the cyclical orientation of discretionary fiscal policy due to the role of automatic stabilisers, which react counter-cyclically by nature. On the empirical side, results depend on modelling choices and datasets, especially the analysed time period.

1.3 Data and methodology

We create a real-time fiscal and macro dataset based on the two detailed annual forecasts of the European Commission, namely the spring and the autumn forecast. We base our analysis on the European Commission forecast because it has the advantage of using the same methodology for all countries in the sample. For some variables, our dataset starts with the forecast vintage of autumn 1997, for others it starts with the vintage of autumn 2002, going up to the forecast vintage of autumn 2020. The sample for which we have both numbers for budget planning and the first outcome covers 21 years in the long sample (1999-2019) and 14 years in the shorter sample (2004-2019). We use the short sample in our baseline analysis and the long sample for robustness checks.

The country panel consists of the 11 original euro area countries excluding Luxembourg[3] (labelled EA11). The dataset includes the main fiscal and macro variables that are necessary to assess the fiscal stance and the cyclical position of the economy: the budget balance (labelled BAL), the cyclically-adjusted budget balance (CAB), the cyclical component of the budget balance (CYC), the primary budget balance (PBAL), public debt (DEBT), nominal and real GDP (GDP), potential GDP (POT), and the output gap (GAP), all in annual frequency. The fiscal variables are expressed in per cent of GDP or potential GDP. The output gap measures the gap between actual and potential GDP, in per cent of potential GDP. Potential GDP is estimated according to the EU's commonly agreed methodology based on a production function approach. The CAB, as measured by the European Commission and used in this chapter, is derived from subtracting the cyclical balance from the total budget balance (Mourre et al., 2004). We calculate interest payments as the difference between the total and the primary balance, and derive the cyclically-adjusted primary balance (CAPB) from subtracting interest payments from the total cyclically-adjusted budget balance. For the CAB, we have data for the long sample period (1999-2019), while the CAPB (as well as PBAL and POT) are only available for the shorter period (2004-2019)

The data stem from the European Commission's forecast webpage for data starting in spring 2013, from Firstrun[4] for the period from 2002 to 2013 and from old forecast documents for the period from 1997 to 2002. We perform extensive checks to ensure the consistency of the overall dataset across the different sources. For the purpose of our analysis, we define the autumn forecast in year *t-1* for *t* as the one which corresponds to information at the time of budget planning (also labelled $E_{t-1}(x_t)$ with x the variable in question), given that the budget plans are elaborated in the autumn of a year for the following year. First realised ex-post outcomes correspond to the data in the spring vintage of year *t+1* for *t* (labelled $x_{t|t+1}$) which corresponds to the first notification of realised values for the year *t*. Finally, the latest vintage of the dataset of autumn 2020 is considered the "final" (no longer revised) estimate of past fiscal and macro variables (labelled $x_{t|T}$).

1.4 Descriptive analysis

1.4.1 Cyclicality of the fiscal stance

In this section, we perform a graphical analysis of the relationship between the economic cycle and the fiscal stance. We use the output gap as measure for the economic cycle. A negative output gap is associated with output below potential, a positive one with output

[3] Austria, Belgium, Germany, Spain, Finland, France, Greece, Ireland, Italy, Netherlands and Portugal. Luxembourg was removed to allow forecast vintages from 2002 onwards, the earliest vintage for Luxembourg would have been 2004, reducing the estimation period by two years.

[4] see http://www.firstrun.eu/

above potential and a closed output gap with output equalling potential. We compare the fiscal stance, measured as the change in the cyclically-adjusted (total or primary) budget balance (CAB and CAPB), with the level of the output gap, i.e. whether a fiscal accommodation or contraction takes place in a situation where the economy is above or below potential. Some papers also use the variation of the output gap as an indicator for the cycle; a concept which shows the acceleration or deceleration of real growth compared to potential growth (see for a discussion Commission, 2004; Golinelli and Momigliano, 2009). We report the comparison of the change in CAB/CAPB with the change in output gap in the Appendix 1.A.

Figure 1.1 plots the variation in the CAB with the level of output gap at the time of budget planning (left-hand panel) and at the time of first outcomes (right-hand panel) for the 2004-2019 period (16 years). The figure shows only a weak link between the two variables, both at the time of budget planning and for first outcomes, which is confirmed by a low correlation coefficient. The graphical analysis hence does not point to a clear pro- or counter-cyclical orientation of fiscal policy, neither during budget planning nor for first outcomes.

Defining four regimes corresponding to the four quadrants of figure 1.1 and counting the observations in each of the quadrants is also informative. We find that fiscal policy is more often pro-cyclical than anti-cyclical in general (more data points in the top-left and bottom-right quadrants than in the other two, see table 1.1). Moreover, fiscal policy is not more pro-cyclical or counter-cyclical ex-ante than ex-post, but we observe more pro-cyclical tightening episodes ex-ante and more pro-cyclical loosening episodes ex-post. We observe especially many planned pro-cyclical tightening episodes during the European Debt crisis.

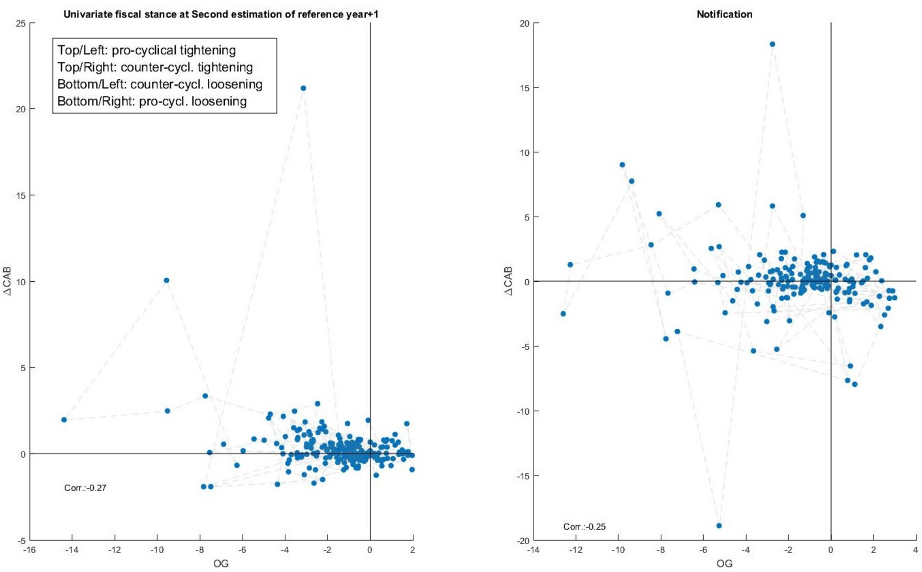


Figure 1.1: Fiscal stance throughout the cycle (ΔCAB and GAP)

Note: a tightening fiscal policy corresponds to an increase in the CAB

	Pro-cycl. tight.	Anti-cycl. tight.	Pro-cycl. loosen.	Anti-cycl. loosen.
$E_{t-1}(x_t)$	81	17	21	57
$x_{t\|t+1}$	72	19	29	56

Table 1.1: Fiscal stance regimes for ΔCAB and GAP

	Pro-cycl. tight.	Anti-cycl. tight.	Pro-cycl. loosen.	Anti-cycl. loosen.
$E_{t-1}(x_t)$	82	12	26	56
$x_{t\|t+1}$	67	13	35	61

Table 1.2: Fiscal stance regimes for $\Delta CAPB$ and GAP

We reach very similar conclusions when analysing the cyclicality of the cyclically-adjusted primary balance (CAPB) instead of the CAB, see figure 1.2 and table 1.2.

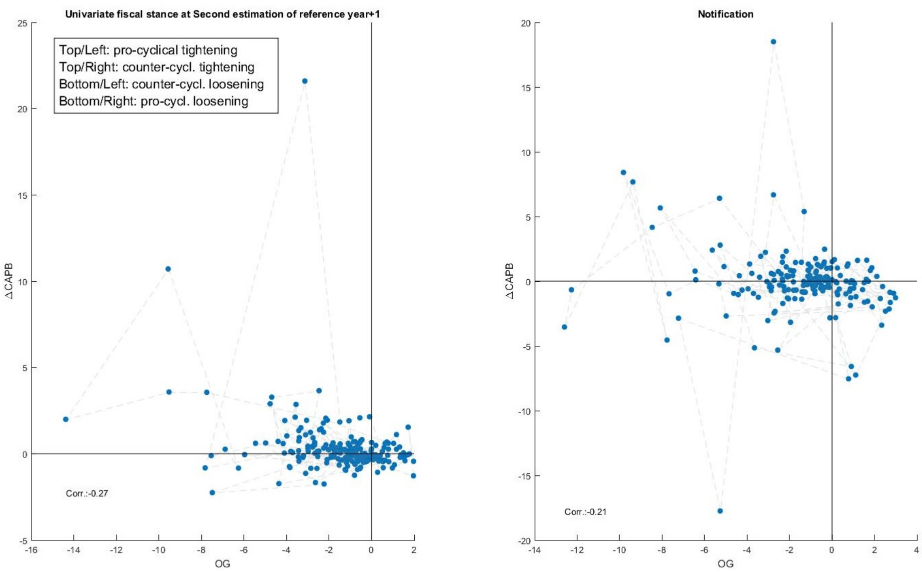


Figure 1.2: Fiscal stance throughout the cycle ($\Delta CAPB$ and GAP)

Note: a tightening fiscal policy corresponds to an increase in the CAPB

1.4.2 Forecast errors for cyclical conditions and the fiscal stance

Since one interest of this chapter is to determine how surprises on cyclical conditions affect the fiscal stance, we analyse in this section the forecast errors (or surprises) of the different variables. The forecast errors are defined as the difference between the one-year ahead forecast for year t made in autumn of year *t-1* and the realised outcome for year t observed in spring of year *t+1*, more precisely we show $\frac{1}{T}\sum_{t=1}^{T}(x_{t|t+1} - E_{t-1}(x_t))$. A positive (respectively, negative) forecast error shows a positive (negative) surprise. Figure 1.3 shows the mean and interquartile range of the forecast errors for the full set of variables.

	GAP	Real GDP growth	POT growth	CAB	CYC	BAL	PBAL	DEBT
RMSE	1.20	1.75	0.69	2.48	0.60	2.57	2.58	6.96
MAE	0.86	1.15	0.46	1.38	0.43	1.45	1.41	4.31
nRMSE	0.53	0.72	0.34	0.89	0.51	0.80	0.94	0.22
nMAE	0.38	0.47	0.22	0.49	0.37	0.45	0.52	0.13

Table 1.3: Dispersion measures for a selection of variables

MAE = mean average error, RMSE = root mean squared error, nMAE = normalized mean average error, nRMSE = normalized root mean squared error

We group these statistics by country and present them per year to give an idea of their temporal evolution. During our sample, the years 2008 and 2009 saw the largest forecast errors for all the variables considered.

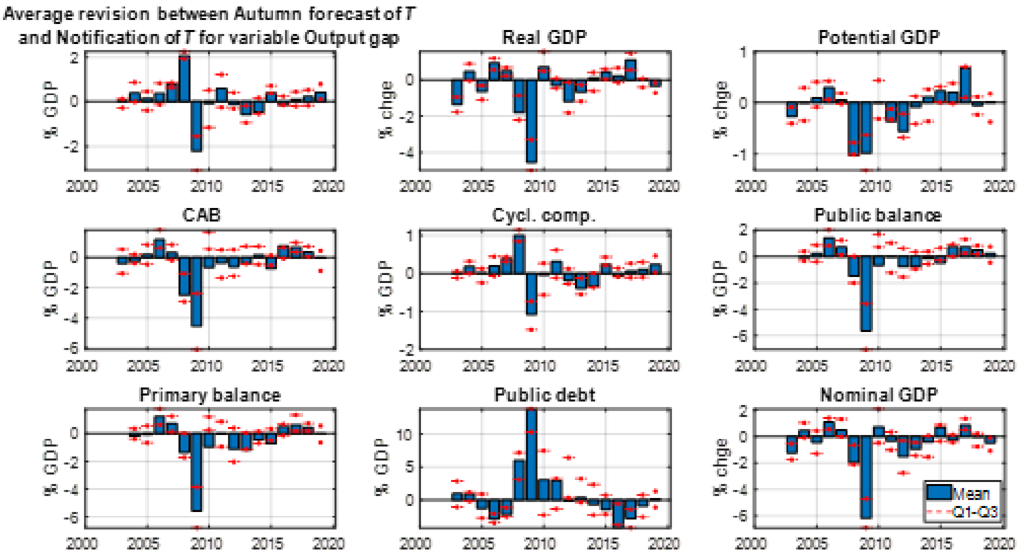


Figure 1.3: Average revision between forecast and outcome for a selection of variables

The largest forecast errors (in relation to the variability of the respective variable) come from measures related to the CAB, BAL and real GDP growth (see different measures of dispersion pooled by year and country in table 1.3). The output gap is somewhat less revised on average than the overall budget balance and the cyclically-adjusted balance. Revisions are the lowest for government debt.

Figure 1.4 resumes the distribution of forecast errors for our two main variables of interest, i.e. the change in the cyclically-adjusted budget balance and the output gap (see Appendix 1.A for the distribution of forecast errors in the change in the cyclically-adjusted primary budget balance). For the output gap, the average forecast error turns out to be positive (error of 0.12 over the full country sample and the 2004-2019 period). The distribution of forecast errors is tilted to the positive side, i.e. we observe more often (106 episodes out of 176) more favorable cyclical conditions than expected. That means that real-time estimates of the output gap tend to overestimate economic downturns and/or underestimate economic upturns. This finding is in line with Larch et al. (2021a), who also show that real-time estimates of the output gap tend to have a pessimistic bias. Fiscal policy is hence inclined to provide too much fiscal support based on real-time output gap estimates. For the change in CAB, we observe a small negative average forecast error (error of -0.20 over the full country sample and the 2004-2019 period), i.e. the cyclically-adjusted balance improves on average less or deteriorates more than expected. This result is however driven by some very large negative forecast errors on the CAB, the distribution is also tilted slightly to the positive side with a higher number of positive surprises (100 episodes out of 176) than negative ones. This result speaks rather against a pro-cyclical bias of fiscal policy orientation in ex-post outcomes compared to ex-ante plans. On the contrary, fiscal outcomes seem more often better than expected.

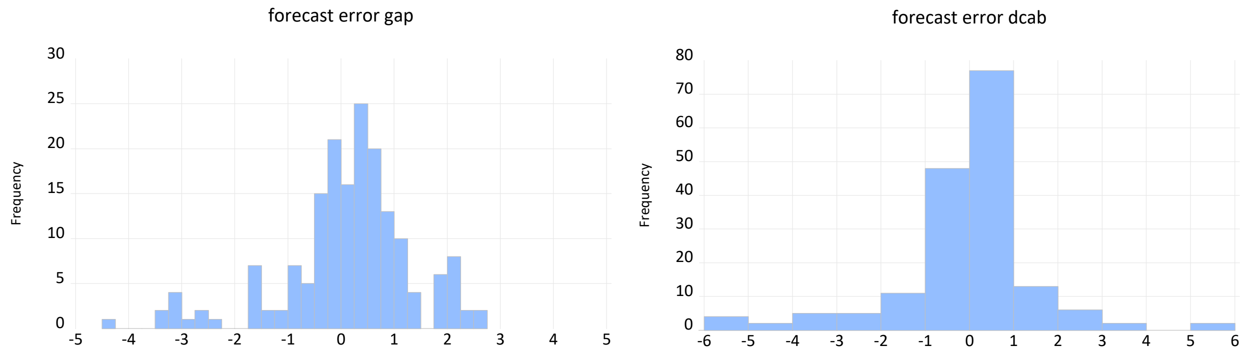


Figure 1.4: Distribution of output gap and cyclically-adjusted balance forecast errors, 2004-2019

Appendix 1.A shows the average forecast error of the forecast error of GAP and ΔCAB for the different countries in the sample.

1.5 Empirical analysis

We now turn to the empirical analysis and estimate the cyclicality of the fiscal stance using a dynamic panel approach. We examine the budget implementation and the outcome phase separately. We perform several robustness checks for the country panel and the time period, paying special attention to the influence of the Great Financial crisis and the European Debt crisis on the results.

1.5.1 Model specification

Following Forni and Momigliano (2005) and Golinelli and Momigliano (2009), we estimate the following fiscal reaction function for the cyclically-adjusted (total or primary) budget balance, both for the budget planning and the outcome phase:

$$\Delta CAB_{i,t} = \alpha + \beta_{CAB} CAB_{i,t-1} + \beta_{GAP} GAP_{i,t} + \beta_{DEBT} DEBT_{i,t-1} + \mu_i + \theta_t + \epsilon_{i,t} \quad (1.1)$$

where the subscript t indicates the time and i the country. $CAB_{i,t}$ stands for the cyclically-adjusted budget balance (or alternatively, $CAPB_{i,t}$ the cyclically-adjusted primary balance). β_{CAB} captures the persistency in the budget balance. $GAP_{i,t}$ is our main variable of interest and shows the cyclical position of the economy. A positive value of β_{GAP} indicates a counter-cyclical stance, while a negative one indicates a pro-cyclical fiscal stance. $DEBT_{i,t-1}$ controls for the level of government debt. A positive value of β_{DEBT} indicates that the higher the initial debt level, the greater the tightening of fiscal policy. μ_i and θ_t are country and time fixed effects. We always include country fixed effects into the model to capture all remaining time-invariant and unobserved country-specific factors. Concerning time fixed effects, which capture common influences of other variables not included in the model, we estimate specifications with and without them[5].

We estimate our baseline model for the 11 euro area countries over the 2004-2019 period (16 observations), using the fixed effect estimator, both for the budget planning and the outcome period. For CAB, we also estimate the model over the long horizon starting from EMU up to 2019 (1999-2019). A possible endogeneity problem might arise from the fact that the cyclical condition $GAP_{i,t}$ is not fully independant from the fiscal stance in a given year t, as government spending is a part of aggregate demand. We employ two empirical strategies to deal with this problem. We use either the lagged value of the output gap $GAP_{i,t-1}$ as a proxy for the output gap in t (the output gaps in t and $t-1$ are highly correlated), which is fully independant of the fiscal policy stance in t, and employ the least square estimator (FE-LS). Alternatively, we use the contemporaneous output gap $GAP_{i,t}$ directly and employ an instrumental variables estimator (FE-IV) controlling for the presence of potentially endogenous variables. Concerning the first strategy, Forni and Momigliano (2005) also give an economic reasoning for using the lagged value of $GAP_{i,t-1}$, suggesting that policy-makers, during the budget planning phase, simply react to current cyclical conditions and not to expected ones. Our results prove robust to both strategies.

In what regards the IV estimator, finding valid instruments for macroeconomic variables is challenging, given the high persistence in the data. Several instruments have been proposed in the literature to control for a possible endogeneity of the output gap, notably lagged values of the output gap and the (GDP-weighted or unweighted) average output gap of the other countries in the sample (see, for example, Beetsma and Giuliodori, 2010; Cimadomo, 2012; Forni and Momigliano, 2005; Pina, 2009). Holm-Hadulla et al. (2010)

[5] A Wald test testing for the joint significance of the time dummies shows that they are jointly significant. We prefer hence the specifications with time fixed effects. We also report the specifications without them to show their influence on the size of the cyclical parameter β_{GAP}.

also use the trend GDP as an instrument. We select five instruments: (i) the nowcast of the output gap, (ii) the forecast of potential output, (iii) the nowcast of potential output, (iv) the forecast of GDP-weighted average gap of the other countries in the sample, and (v) the nowcast of GDP-weighted average gap of the other countries in the sample. To check the relevance of our instruments, we run auxiliary regressions for the potential endogenous output gap variable. We also test for weak instruments using the F-test and perform a Sargan test of the overidentifiying restrictions.

Our panel model includes the lagged value of the dependant variable $CAB_{i,t-1}$, which induces a potential bias into the fixed effet estimator, also called "Nickell"-bias (Nickell, 1981). The "Nickell" bias is decreasing with the number of time observations T. We argue that the bias is relatively small as the number of time observation is large compared to the number of cross sections N in our model, in particular in the long panel ($T > 20$). Also, it has been shown that alternative estimators, notably the GMM estimator, would not alleviate the problem in our case where $T > N$ (see also Judson and Owen, 1999 and Checherita-Westphal and Zdarek, 2017).

We check for cross-sectional dependance in the residuals, using the Breush-Pagan LM test (Breusch and Pagan, 1980)[6]. It indicates some cross-sectional dependence in the residuals in the model, both for the budget planning and the first outcome phase. Hence, we cluster standard errors at the period level using the cross section SUR Panel Corrected Standard Error (PCSE). This setting controls for contemporaneous correlations between the residuals for cross-sections. Having in mind that this gives conservative values for standard errors, we also comment, when appropriate, the results with ordinary standard errors.

The variables in the specification for each, the budget planning and the first outcome, all come from the same vintage in order to assure a common information set and avoid eventual problems due to methodological changes. For example, for the planning phase, the dependant variable is defined as the first difference of the expected value of CAB in $t-1$ for t and the expected value of CAB in $t-1$ for $t-1$. The lagged independant variables of CAB, GAP and DEBT represent the expected values in $t-1$ for $t-1$. The output gap in the FE-IV is the expected gap in $t-1$ for t. All variables hence come from the same vintage of $t-1$ during the budget planning phase (precisely, the autumn forecast of the European Commission).

1.5.2 Main empirical results

Table 1.4 provides estimates of our different specifications for $\Delta CAB_{i,t}$. Columns (1) to (3) show results of ex ante fiscal reaction functions, while columns (4) to (6) provide estimates of the fiscal reaction functions with first outcome data. In the specifications with the lagged output gap (columns (1) and (2) of table 1.4), the coefficient of the cyclical

[6] The Breush-Pagan LM test is more suitable than the Pesaran CD test (Pesaran, 2004) in this case, because T>N.

	Budget planning			First outcome		
	(1)	(2)	(3)	(4)	(5)	(6)
Method	LS-FE	LS-FE	IV-FE	LS-FE	LS-FE	IV-FE
$CAB_{i,t-1}$	-0.416***	-0.512***	-0.303***	-0.297**	-0.348**	-0.122
	[-5.929]	[-7.078]	[-2.683]	[-2.399]	[-2.523]	[-0.545]
$GAP_{i,t-1}$	-0.109	-0.240***		-0.026	0.056	
	[-1.351]	[-2.921]		[-0.194]	[0.291]	
$GAP_{i,t}$			-0.209*			-0.104
			[-1.800]			[-0.619]
$DEBT_{i,t-1}$	0.003	-0.003	0.009	0.024	0.029	-0.032
	[0.240]	[-0.263]	[0.454]	[1.070]	[0.996]	[-1.068]
Constant	-0.928	-0.858	-1.354	-2.587	-3.014	2.485
	[-1.029]	[-0.879]	[-0.909]	[-1.377]	[-1.280]	[0.995]
Nb. of obs.	176	176	176	176	176	176
Adj. R^2	0.539	0.699	0.579	0.133	0.257	0.088
Country FE	yes	yes	yes	yes	yes	yes
Time FE	no	yes	yes	no	yes	yes

Table 1.4: Fiscal reaction functions in the euro area (2004-2019) for $\Delta CAB_{i,t}$

Asterisks *, ** and *** denote significance at 10%, 5% and 1% significance levels, respectively. T-statistics are given in parentheses. Robust standard errors (cross-section SUR PCSE). Country and time fixed effects are not reported. Equation (6) includes only the lagged output gap and the forecast and nowcast of the lagged output gap of other countries as instruments, as the potential POT is not significant in the first-stage auxiliary regression.

condition is negative, indicating a pro-cyclical orientation of fiscal policy during budget planning. The coefficient is significant in the specification with time fixed effects[7]. The estimated coefficient in (2) implies that a (positive) output gap of 1 per cent leads to a worsening of the CAB by 0.24 per cent of GDP on average for euro area countries. The coefficient of the lagged CAB has the expected negative sign and is highly significant. The initial debt level does not have a significant impact on the fiscal policy orientation during the budget planning phase. The results are similar when analyzing the different specifications for $\Delta CAPB_{i,t}$ (see table 1.5). The coefficients of the lagged output gap (columns 1 and 2 of table 1.5) are negative and significant in what regards the specification with time fixed effects: an output gap of 1 per cent reduces the CAPB by about 0.19 per cent of GDP on average for euro area countries, implying a pro-cyclical fiscal policy stance. The lagged dependant variable has a significant and negative influence on the fiscal stance and the debt level a not significant one.

The IV specifications broadly confirm these results (columns (3) of respectively table 1.4 and table 1.5). The contemporanous output gap has a significant negative sign, in-

[7]The coefficient is also significant in the specification using only country fixed effects with ordinary standard errors. The size of the estimated coefficient is slightly smaller than in the specification with country and time fixed effects.

dicating a pro-cyclical stance of fiscal policy. The coefficient is close in size to the OLS specifications, the significance is however slightly reduced. To check that our instruments are strong, we estimate the so-called first-stage regression (*i.e.* we regress the potentially endogeneous variable $GAP_{i,t}$ on the selected set of instruments and other exogenous variables). Table 1.6 shows the results of a FE-LS estimate for the first-stage regression with time and country fixed effects. The estimated coefficients on all instruments are statistically different from 0, we hence conclude that they are relevant and strong[8]. Then we perform a Sargan test of the overidentifiying restrictions when the number of instruments exceeds the number of endogenous regressors, which is the case for the IV-FE estimate at the budget planning stage[9].

Turning to the ex post realisation, we obtain either a small negative (column (4) of table 1.4) or a small positive coefficient (column (5) of table 1.4) on the output gap, which is not significant in any of the specifications[10]. We hence conclude on a broadly a-cyclical fiscal policy orientation in outcome data. The results are similar for $\Delta CAPB_{i,t}$ (columns (4) and (5) of table 1.5). The coefficients of the lagged CAB/CAPB are close to those with real-time data and highly significant. The initial debt level, insignificant in three of the four specifications, shows a small significant positive impact on CAPB in the specification with country and time fixed effects (column (5) of table 1.5). A debt level of 1 per cent of GDP indicates a fiscal tightening of the cyclically-adjusted primary balance of 0.04 per cent of GDP. The fit of the ex post regressions is generally worse than the one using real-time information. We conclude that the expected orientation of fiscal policy can be relatively well explained with cyclical conditions and starting levels of budget balances, while ex post outcomes are less well explained by our model. We also examine to what extent further revisions affect these ex post results and estimate the model with the latest available "final" vintage of data (more specifically, the autumn 2020 forecast). These results attest an a-cyclical (not significant) ex post fiscal orientation both for ΔCAB and $\Delta CAPB$, see Appendix 1.B.

We reach the same conclusions if we look at the change instead of the level of the output gap. Fiscal policy has been pro-cyclical in the planning phase during cyclical downturns or upswings, i.e. showing an opposite sign to a change of a given sign of the output gap, and a-cyclical (not significant) in first outcomes, see Appendix 1.B.

To conclude, our baseline results show that fiscal policy in the euro area has on average been pro-cyclical at the planning stage over the analysed period and that this pro-cyclical orientation does not result from a distorted or incorrect perception of the cyclical situation. On the contrary, the actual results are better from a fiscal stabilisation perspective than planned ones, pointing to an a-cyclical orientation of fiscal policy.

[8]The F-stat of the joint nullity test is 232.756

[9]The p-value of the Sargan test is 0.125 for the IV-FE specification for ΔCAB at the budget planning stage (column (3) of table 1.4) and 0.177 for the IV-FE specification for $\Delta CAPB$ at the budget planning stage (column (3) of table 1.5), meaning that the overidentifying restrictions are valid.

[10]The coefficients are also not significant when using ordinary standard errors

These results contrast somewhat with earlier studies on real-time data, which predominantly find a counter-cyclical orientation of fiscal policy in real-time (see Cimadomo, 2012; Forni and Momigliano, 2005; Golinelli and Momigliano, 2009; Pina, 2009). They are in line with these studies for ex-post outcomes, which also predominantly show an a-cyclical (non significant) fiscal policy orientation. This difference could be explained by several factors, the most important being the different dataset (most of the earlier studies, except for Pina (2009), use OECD data instead of Commission data) and the different time period (most of the earlier studies use data up to the early 2000s). To test this hypothesis, we estimate the CAB specification[11] (2) of table 1.4 from the earliest date available (1997) and, alternatively, from the start of EMU (1999) up to the GFC (2007). The coefficient of the output gap turns indeed very slightly positive, but it is not significant in any of the specifications. We hence cannot confirm a significant counter-cyclical fiscal policy in real-time on an earlier time sample with our data[12]. Our results confirm however the findings from Aldama and Creel (2022), who also find a pro-cyclical fiscal policy in real-time in 15 EU-countries over the period 1995-2017. They do not perform ex post estimations. More surprisingly is the difference of our results with those of Gootjes and de Haan (2022), who also use Commission data and a similar country (in the specification for euro area countries) and time sample (2000-2015), finding a counter-cyclical fiscal policy ex ante.

Also, our result shows the average reaction for the 11 euro area countries included in the sample. Some countries show a more pro-cyclical fiscal stance at the budget planning stage than others which have a more counter-cyclical fiscal stance (see section 1.5.3). Including or excluding these countries from our sample would drive the average outcome in one direction or the other.[13]

1.5.3 Extensions and robustness checks

Results since the start of EMU

As a first robustness check, we estimate the equations for the change in CAB over the long horizon since the start of EMU (1999-2019), see table 1.7. In addition to showing the fiscal orientation in the euro area since the start of EMU, this also provides a robustness check for the "Nickell"-bias, which decreases with the number of time series observations T, which amounts to above 20 in the long sample (see Judson and Owen, 1999).

The results show a significant pro-cyclical fiscal orientation in real-time since the start

[11] The cited studies use mostly the cyclically-adjusted primary balance CAPB as an indicator for the fiscal stance, for which we have however only data starting in 2004.

[12] We do not report results here given the short estimation period which renders results fragile.

[13] In particular, Greece and Ireland are the countries with the strongest fiscal reaction to the output gap according to our individual regressions. These countries are also those where the output gap and the change in CAB (or CAPB) have larger amplitudes. When removing these two euro area countries from our sample, we find an a-cyclical fiscal stance in the other 9 countries, both at the planning phase and with first outcomes. This is not driven just by the European Debt crisis, which we show in section 1.5.3.

	Budget planning			First outcome		
	(1)	(2)	(3)	(4)	(5)	(6)
Method	LS-FE	LS-FE	IV-FE	LS-FE	LS-FE	IV-FE
$CAPB_{i,t-1}$	-0.482***	-0.562***	-0.423***	-0.351***	-0.400***	-0.219
	[-7.041]	[-8.250]	[-3.471]	[-2.738]	[-6.127]	[-0.767]
$GAP_{i,t-1}$	-0.124	-0.186**		-0.020	0.112	
	[-1.540]	[-2.296]		[-0.147]	[1.072]	
$GAP_{i,t}$			-0.219*			-0.120
			[-1.907]			[-0.702]
$DEBT_{i,t-1}$	0.007	0.017	0.014	0.024	0.042***	-0.036
	[0.634]	[1.442]	[0.744]	[1.074]	[2.795]	[-1.151]
Constant	-0.083	-1.015	-0.821	-1.802	-3.107**	3.086
	[-0.090]	[-1.065]	[-0.541]	[-0.984]	[-2.537]	[1.198]
Nb. of obs.	165	165	165	176	176	176
Adj. R^2	0.610	0.733	0.696	0.159	0.303	0.136
Country FE	yes	yes	yes	yes	yes	yes
Time FE	no	yes	yes	no	yes	yes

Table 1.5: Fiscal reaction functions in the euro area (2004-2019) for $\Delta CAPB_{i,t}$

Asterisks *, ** and *** denote significance at 10%, 5% and 1% significance levels, respectively. T-statistics are given in parentheses. Robust standard errors (cross-section SUR PCSE). Country and time fixed effects are not reported. Equation (6) includes only the lagged output gap and the forecast and nowcast of the lagged output gap of other countries as instruments, as the potential POT is not significant in the first-stage auxiliary regression.

Method	LS-FE
$GAP_{i,t-1}$	0.867***
	[18.344]
$POT_{i,t}$	0.351**
	[2.181]
$POT_{i,t-1}$	-0.360**
	[-2.600]
$OTHER\_GAP_{i,t}$	-6.505***
	[-6.973]
$OTHER\_GAP_{i,t-1}$	5.247***
	[6.351]
$CAB_{i,t-1}$	-0.006
	[-0.242]
$DEBT_{i,t-1}$	0.007
	[1.315]
Constant	-1.822**
	[-2.178]
Nb. of obs.	176
Adj. R^2	0.936
Country FE	yes
Time FE	yes

Table 1.6: First-stage regression in real-time

Asterisks *, ** and *** denote significance at 10%, 5% and 1% significance levels, respectively. T-statistics are given in parentheses. Ordinary standard errors. Country and time fixed effects are not reported.

	Budget planning			First outcome		
	(1)	(2)	(3)	(4)	(5)	(6)
Method	LS-FE	LS-FE	IV-FE	LS-FE	LS-FE	IV-FE
$CAB_{i,t-1}$	-0.386***	-0.457***	0.0545	-0.266**	-0.323***	0.066
	[-6.304]	[-7.040]	[0.380]	[-2.578]	[-2.903]	[0.281]
$GAP_{i,t-1}$	-0.042	-0.153**		0.016	0.089	
	[-0.586]	[-1.989]		[0.141]	[0.605]	
$GAP_{i,t}$			-0.169*			-0.079
			[-1.632]			[-0.528]
$DEBT_{i,t-1}$	0.003	0.002	0.001	0.020	0.025	-0.035
	[0.341]	[0.189]	[0.081]	[1.124]	[1.281]	[-1.095]
Constant	-0.665	-0.830	0.124	-1.893	-2.372	3.064
	[-0.944]	[-1.139]	[0.086]	[-1.345]	[-1.538]	[1.192]
Nb. of obs.	231	231	231	231	231	231
Adj. R^2	0.497	0.618	0.040	0.111	0.249	-0.086
Country FE	yes	yes	yes	yes	yes	yes
Time FE	no	yes	yes	no	yes	yes

Table 1.7: Fiscal reaction functions in the euro area (1999-2019) for $\Delta CAB_{i,t}$

Asterisks *, ** and *** denote significance at 10%, 5% and 1% significance levels, respectively. T-statistics are given in parentheses. Robust standard errors (cross-section SUR PCSE). Country and time fixed effects are not reported. Equation (3) and (6) include the lagged output gap and the forecast and nowcast of the lagged output gap of other countries as instruments, as the potential POT is not available over the long period.

of EMU, in what regards the specification with country and time fixed effects. The coefficient in (2) means that a (positive) output gap of 1 per cent leads to a worsening of the CAB by 0.15 per cent of GDP, compared to a worsening of the CAB of 0.24 in the shorter sample. The coefficient of the other variables are also comparable in size and in significance to the ones in the shorter sample.[14] The coefficient on the output gap in outcome data is slightly positive but never significant. We hence also conclude on a broadly a-cyclical stance in fiscal outcomes over the long period since 1999.

Impact of the crisis periods

In order to analyse the influence of the crisis years on the above results, we perform two exercises: In a first step, we estimate the above specifications with the lagged output gap with time dummies for the crisis years[15], namely $dummy_{GFC} = 1$ for the years 2008 and 2009 and $dummy_{EDC} = 1$ for the years 2012 and 2013[16], and 0 otherwise. Column (1) of table 1.8 shows that adding time dummies for the crisis periods does not alter the conclusions of the baseline results, the coefficient of the output gap remains negative[17] and close in size to table 1.4 for the budget planning phase. The crises dummies reveal a planned fiscal tightening during the GFC and a small planned fiscal loosening during the EDC, both coefficients being however not significant. In a second step, we verify the influence of the crisis years on the conclusion on the cyclicality of the fiscal stance by interacting the time dummies with the output gap. Column (2) of table 1.8 confirms the pro-cyclicality of the fiscal stance during the budget planning period outside of the crisis years. The coefficient indicates that a (positive) output gap of 1 per cent leads to a worsening of the CAB by 0.30 per cent of GDP during non-crisis years, which is even slightly higher in absolute terms than in our baseline results. During the GFC, the coefficient of the output gap shows a non-significant counter-cyclical stance, while we observe a weakly significant pro-cyclical fiscal stance (tightening) during the EDC. We conclude hence that the baseline findings of a pro-cyclical fiscal orientation in real-time are not driven by the two crisis periods[18]. The fiscal orientation during the EDC was however also clearly oriented in a pro-cyclical manner. These results are confirmed when performing the same analysis over the long time sample (see Appendix 1.B).

In what regards the first outcomes, column (3) of table 1.8 confirms an a-cyclical orientation of fiscal policy when controlling for the two crisis periods. The crisis dummies

[14] We don't comment the IV-specification in table 1.7, given that the coefficients of the lagged CAB variable have an unexpected sign. There are shown for the purpose of completness only.

[15] The specifications with time fixed effects already control for time specific factors. The interest in using specific time dummies lies in testing the influence of the crises years jointly.

[16] We define common dummies for the full country set. There are set 1 for years where real GDP growth was negative in the euro area as a whole, and 0 otherwise.

[17] The coefficient is not significant in the estimation with robust standard errors and strongly significant with ordinary standard errors.

[18] The conclusions are the same and coefficients very close to table 1.8 when using the contemporanous output gap instead of the lagged one

	Budget planning		First outcome	
	(1)	(2)	(3)	(4)
Method	LS-FE	LS-FE	LS-FE	LS-FE
$CAB_{i,t-1}$	-0.423***	-0.516***	-0.308***	-0.361**
	[-5.897]	[-7.334]	[-2.691]	[-2.536]
$GAP_{i,t-1}$	-0.123		0.115	
	[-1.541]		[0.858]	
$DEBT_{i,t-1}$	0.006	-0.007	0.019	0.027
	[0.515]	[-0.540]	[0.974]	[0.848]
$dummy_{GFC}$	0.566		-2.580***	
	[0.938]		[-3.099]	
$dummy_{EDC}$	-0.135		-0.135	
	[-0.242]		[-0.197]	
$GAP_{i,t-1} * dummy_{GFC}$		0.392		0.081
		[1.008]		[0.113]
$GAP_{i,t-1} * dummy_{EDC}$		-0.196*		0.227
		[-1.975]		[0.873]
$GAP_{i,t-1} * (1 - dummy_{GFC} - dummy_{EDC})$		-0.295***		-0.020
		[-3.229]		[-0.090]
Constant	-1.295	-0.663	-1.642	-2.804
	[-1.331]	[-0.704]	[-0.996]	[-1.186]
Nb. of obs.	176	176	176	176
Adj. R^2	0.543	0.678	0.205	0.256
Country FE	yes	yes	yes	yes
Time FE	no	yes	no	yes

Table 1.8: Fiscal reaction functions in the euro area (2004-2019) for $\Delta CAB_{i,t}$ controlling for crisis years

Asterisks *, ** and *** denote significance at 10%, 5% and 1% significance levels, respectively. T-statistics are given in parentheses. Robust standard errors (cross-section SUR PCSE). Country and time fixed effects are not reported.

	Budget planning		First outcome	
	(1)	(2)	(3)	(4)
Method	LS-FE	LS-FE	LS-FE	LS-FE
$CAPB_{i,t-1}$	-0.492***	-0.566***	-0.357***	-0.408***
	[-6.878]	[-8.562]	[-3.022]	[-2.896]
$GAP_{i,t-1}$	-0.134*		0.117	
	[-1.682]		[0.865]	
$DEBT_{i,t-1}$	0.012	0.014	0.019	0.039
	[0.983]	[1.197]	[0.966]	[1.304]
$dummy_{GFC}$	0.827		-2.479**	
	[1.309]		[-2.798]	
$dummy_{EDC}$	0.310		-0.025	
	[0.537]		[-0.034]	
$GAP_{i,t-1} * dummy_{GFC}$		0.467		0.126
		[1.384]		[0.184]
$GAP_{i,t-1} * dummy_{EDC}$		-0.225**		0.330
		[-2.044]		[1.325]
$GAP_{i,t-1} * (1 - dummy_{GFC} - dummy_{EDC})$		-0.221**		0.023
		[-2.496]		[0.111]
Constant	-0.634	-0.805	-0.861	-2.856
	[-0.630]	[-0.871]	[-0.535]	[-1.241]
Nb. of obs.	165	165	176	176
Adj. R^2	0.623	0.740	0.224	0.308
Country FE	yes	yes	yes	yes
Time FE	no	yes	no	yes

Table 1.9: Fiscal reaction functions in the euro area (2004-2019) for $\Delta CAPB_{i,t}$ controlling for crisis years

Asterisks *, ** and *** denote significance at 10%, 5% and 1% significance levels, respectively. T-statistics are given in parentheses. Robust standard errors (cross-section SUR PCSE). Country and time fixed effects are not reported.

reveal a sizeable and significant fiscal loosening during the GFC and a small but non-significant loosening during the EDC. The conclusions on the cyclicality of fiscal policy are not altered when considering only non-crisis years (see column (4) of table 1.8): we observe a small but non significant coefficient on the output gap, indicating a broadly a-cyclical fiscal stance for first fiscal outcomes. During the crisis years, the coefficient of the output gap is slightly positive but also non significant. Ex-post we hence do not confirm the pro-cyclical nature of fiscal policy during the EDC. See Appendix 1.B for the estimations over the long time sample.

Table 1.9 presents the replication of the analysis for $\Delta CAPB_{i,t}$, confirming largely the above findings for $\Delta CAB_{i,t}$.

Group	Country	Budget planning	First outcome
Group 1 (- sign on $GAP_{i,t-1}$ in budget planning)	GR	-0.449**	-0.034
		[-2.230]	[-0.086]
	FR	-0.337**	-0.348*
		[-2.335]	[-1.981]
	IT	-0.145	-0.013
		[-0.943]	[-0.136]
	ES	-0.086	-0.327
		[-1.027]	[-1.463]
	FI	-0.040	-0.106
		[-0.226]	[-1.340]
	PT	-0.004	0.033
		[-0.032]	[0.126]
Group 2 (+ sign on $GAP_{i,t-1}$ in budget planning)	BE	0.096	-0.462*
		[0.599]	[-1.945]
	NL	0.134	-0.240
		[0.621]	[-1.161]
	AT	0.315*	0.076
		[2.049]	[0.569]
	DE	0.521***	0.158
		[3.956]	[1.393]
	IE	1.129***	1.067
		[4.323]	[1.429]

Table 1.10: Individual fiscal reaction functions (2004-2019) for $\Delta CAB_{i,t}$: coefficient on lagged output gap

Asterisks *, ** and *** denote significance at 10%, 5% and 1% significance levels, respectively. T-statistics are given in parentheses. Ordinary least squares. Ordinary standard errors.

Heterogeneity inside the euro area

In this section we look deeper in the question whether Euro area countries or groups of countries differ in terms of their fiscal stance. We estimate our fiscal reaction functions for ΔCAB with the lagged output gap for each country individually by OLS. Then we group countries according to their ex ante fiscal stance. In a first group we gather the countries whose fiscal policy is pro-cyclical ex ante (or a-cyclical with a negative coefficient on the lagged output gap), whereas the other group contains countries with a counter-cyclical ex ante fiscal stance (or a-cyclical with a positive coefficient on the lagged output gap). The two groups are presented in table 1.10 with the coefficients on the output gap obtained by an OLS estimate of the fiscal reaction function with ex ante information (column 1) and first outcome data (column 2).

The group of countries with an ex ante pro-cyclical fiscal stance is made up of Greece, France, Italy, Spain, Finland and Portugal. Apart from France and Finland, these coun-

Group of countries	Group 1 (pro-cyclical)		Group 2 (counter-cyclical)	
Data vintage	Budget planning	First outcome	Budget planning	First outcome
Method	LS-FE	LS-FE	LS-FE	LS-FE
$CAB_{i,t-1}$	-0.359***	-0.370***	-0.805***	-0.537*
	[-4.627]	[-2.904]	[-9.810]	[-1.766]
$GAP_{i,t-1}$	-0.213**	0.001	0.642***	1.002
	[-2.375]	[0.006]	[2.912]	[1.090]
$DEBT_{i,t-1}$	0.012	0.061**	-0.043**	0.025
	[1.032]	[2.193]	[-2.613]	[0.370]
Constant	-2.053*	-6.466**	2.511**	-2.264
	[-1.846]	[-2.407]	[2.107]	[-0.479]
Nb. of obs.	96	96	80	80
Adj. R^2	0.550	0.376	0.841	0.257
Country FE	yes	yes	yes	yes
Time FE	yes	yes	yes	yes

Table 1.11: Fiscal reaction functions in subgroups of euro area countries (2004-2019) for $\Delta CAB_{i,t}$

Asterisks *, ** and *** denote significance at 10%, 5% and 1% significance levels, respectively. T-statistics are given in parentheses. Robust standard errors (cross-section SUR PCSE). Country and time fixed effects are not reported.

tries were those most affected by the EDC. France also had a deficit above the 3 % threshold during this period and might have run a pro-cyclical fiscal policy in order not to aggrevate its fiscal position. On the other side, Belgium, the Netherlands, Austria, Germany and Ireland have a planned counter-cyclical fiscal stance. The euro area estimate presents the average results of the 11 individual estimates. Hence, our result of an ex ante pro-cyclical fiscal stance in the euro area countries is driven mostly by Greece, France, and to a lesser extent Italy, Spain, Finland and Portugal, while the other countries, and in particular Ireland and Germany, temper this result.

Most countries keep the same fiscal orientation (*i.e.* the sign of the coefficient on lagged output gap) in the first release as planned in the budget, with the exception of the Netherlands (whose coefficients are not significant), Portugal (whose coefficients are very low and not significant) and Belgium. Except for France and Belgium, the coefficient is not significant in the first outcome for none of the countries.

We then estimate the panel specification (2) (with the lagged output gap and country and time fixed effects) for each of the identified groups of countries and for the two vintages of data: budget planning and first outcome, see table 1.11[19]. Again, the ex ante estimates explain more variance than the ex post estimates.

[19]We show here the results for the split country sample. We also perform the estimation with the full country sample, using interaction effects with a dummy for the ex-ante pro-cyclical country group instead, see 1.B.4 in appendix 1.B. This does not alter the coefficient estimates.

Countries that are pro-cyclical ex ante become a-cyclical (with a positive but non significant coefficient) ex post. The coefficient on debt, which is not significant ex ante, becomes significant ex post. It is positive, indicating a tendency for the fiscal balance to improve as the debt increases. Countries that have a counter-cyclical fiscal policy ex ante also are a-cyclical ex post (the coefficient on the lagged output gap becomes insignificant, although it increases relative to the ex ante estimate). While the coefficient on debt is negative ex ante, indicating an inverse relationship between the debt level and the fiscal orientation, it becomes non significant ex post.

We obtain very similar results when we estimate individual fiscal reaction functions for $\Delta CAPB$ instead of ΔCAB (see tables 1.B.5 and 1.B.6 at the end of appendix 1.B).

Tightening and loosening episodes

An important question is what could explain the intention of governments to implement a pro-cyclical fiscal policy, which is sub-optimal from a macroeconomic stabilisation point of view. Two possible explanations proposed in the introductory section are a general deficit bias of fiscal policy or forced fiscal consolidation during times of market stress. The first would rather be in line with pro-cyclical loosening episodes while the later would rather occur during tightening episodes. We hence examine the cyclicality during tightening and loosening episodes seperately in this section.

We run our baseline regressions with and without time fixed effects with the FE-LS estimator, controlling for episodes of tightening and loosening, on the full sample of countries from 2004 to 2019. In concrete terms, we interact the output gap with a dummy for tightening episodes and a dummy for loosening episodes. Results are shown in table 1.12 for the change in CAB and in table 1.13 for the change in CAPB.

We find that the fiscal stance is pro cyclical ex ante during tightening episodes, and rather a-cyclical during loosening episodes at the budget planning phase, both for the change in CAB and the change in CAPB (see columns (2) of table 1.12 and table 1.13). In our prefered specification with time fixed effects (see columns (4) of table 1.12 and table 1.13), the fiscal stance is a-cyclical ex post during tightening episodes and either a-cyclical (for the change in CAB) or even counter-cyclical (for the change in CAPB) during loosening episodes. We hence argue that the pro-cylicality of ex ante fiscal policy during tightening episodes is related to forced consolidation due to market stress or the intention of governments to signal a "stronger" tightening in their fiscal balances in difficult economic times compared to what is really implemented in the end.

1.6 Conclusions

In this chapter, we analyse the cyclical orientation of fiscal policy in the euro area since 1999 and 2004 until 2019 during budget planning and for first outcomes, using a new real-time dataset based on the European Commission's biannual forecasts. Our empirical

	Budget planning		First outcome	
	(1)	(2)	(3)	(4)
Method	LS-FE	LS-FE	LS-FE	LS-FE
$CAB_{i,t-1}$	-0.384***	-0.492***	-0.206*	-0.260*
	[-5.855]	[-6.726]	[-1.679]	[-1.882]
$GAP_{i,t-1} * (\Delta CAB_{i,t} > 0)$	-0.230***	-0.306***	-0.345**	-0.243
	[-2.794]	[-3.711]	[-2.448]	[-1.329]
$GAP_{i,t-1} * (\Delta CAB_{i,t} < 0)$	0.118	-0.089	0.218	0.303
	[1.133]	[-0.811]	[1.551]	[1.538]
$DEBT_{i,t-1}$	0.001	-0.004	0.026	0.030
	[0.098]	[-0.364]	[1.234]	[1.087]
Constant	-0.706	-0.696	-2.729	-3.049
	[-0.897]	[-0.726]	[-1.534]	[-1.365]
Nb. of obs.	176	176	176	176
Adj. R^2	0.596	0.686	0.235	0.356
Country FE	yes	yes	yes	yes
Time FE	no	yes	no	yes

Table 1.12: Fiscal reaction functions in the euro area (2004-2019) for $\Delta CAB_{i,t}$ controlling for tightening and loosening episodes

Asterisks *, ** and *** denote significance at 10%, 5% and 1% significance levels, respectively. T-statistics are given in parentheses. Robust standard errors (cross-section SUR PCSE). Country and time fixed effects are not reported.

	Budget planning		First outcome	
	(1)	(2)	(3)	(4)
Method	LS-FE	LS-FE	LS-FE	LS-FE
$CAPB_{i,t-1}$	-0.442***	-0.539***	-0.240*	-0.289**
	[-6.843]	[-7.681]	[-1.943]	[-2.138]
$GAP_{i,t-1} * (\Delta CAPB_{i,t} > 0)$	-0.249***	-0.264***	-0.418***	-0.270
	[-3.028]	[-3.206]	[-3.038]	[-1.574]
$GAP_{i,t-1} * (\Delta CAPB_{i,t} < 0)$	0.086	-0.065	0.237*	0.366**
	[0.868]	[-0.655]	[1.733]	[2.101]
$DEBT_{i,t-1}$	0.006	0.017	0.028	0.044*
	[0.614]	[1.389]	[1.323]	[1.660]
Constant	-0.032	-0.987	-2.354	-3.503*
	[-0.041]	[-1.044]	[-1.367]	[-1.668]
Nb. of obs.	165	165	176	176
Adj. R^2	0.661	0.748	0.291	0.430
Country FE	yes	yes	yes	yes
Time FE	no	yes	no	yes

Table 1.13: Fiscal reaction functions in the euro area (2004-2019) for $\Delta CAPB_{i,t}$ controlling for tightening and loosening episodes

Asterisks *, ** and *** denote significance at 10%, 5% and 1% significance levels, respectively. T-statistics are given in parentheses. Robust standard errors (cross-section SUR PCSE). Country and time fixed effects are not reported.

results indicate a pro-cyclical behaviour of fiscal plans on average for the euro area, for both the cyclically-adjusted total balance and the cyclically-adjusted primary balance. These results are robust to different specifications and estimation methods. In our preferred baseline specification over the 2004-2019 period, a (positive) output gap of 1 per cent worsens the cyclically-adjusted balance by about -0.24 per cent and the cyclically-adjusted primary balance by about -0.19 per cent during budget planning. Over the longer 1999-2019 period, a (positive) output gap of 1 per cent is associated with a worsening of the cyclically-adjusted balance by about -0.15 per cent. Turning to first outturns, our results suggest a broadly a-cyclical behaviour of fiscal policy in the euro area, with non-significant coefficients of the output gap in the estimations, both for the longer and the shorter period. We do not find any evidence for an impact of debt levels on the fiscal policy orientation, neither during fiscal planning nor in first outcomes.

These findings are confirmed when controlling for the Great Financial crisis or the European Debt crisis, which occurred during the sample period. Fiscal plans were also pro-cyclical excluding these crisis years, and, according to our empirical results, counter-cyclical (but not significant) during the Great Financial crisis and significantly pro-cyclical during the European Debt crisis. Our results contrast somewhat with earlier studies on the fiscal policy orientation in real-time, which find that fiscal plans were conceived counter-cyclically and that surprises on the economic conditions led to a broadly a-cyclical orientation ex-post (notably Cimadomo, 2012; Forni and Momigliano, 2005; Golinelli and Momigliano, 2009; Pina, 2009), given an optimistic bias on the output gap estimate in real-time. Our findings are in line with the findings of a more recent paper by Aldama and Creel (2022), who also find a pro-cyclical fiscal orientation during budget planning in OECD countries (covering a large panel of euro area countries). We also show that the fiscal orientation is not uniform in the euro area that the average result is driven by some countries with a strong pro-cyclical ex ante fiscal policy orientation. We observe a more pro-cyclical behaviour of fiscal plans for countries such as Greece, France, Italy, Spain, and Finland and a more counter-cyclical behaviour in fiscal plans for a group of countries covering Ireland, Germany, Austria, and to a lesser extent also the Netherlands and Belgium. The heterogeneity of the cyclicality of fiscal policy is very important in the euro area.

Our results imply that, on average for euro area countries, the cyclical orientation of fiscal policy is already suboptimal from a macroeconomic stabilisation perspective during budget planning and it is not just the misjudgement on cyclical conditions that drives the result of an a- or pro-cyclical behaviour in fiscal outturns in the euro area. We also find an optimistic bias on cyclical conditions in real-time, which might motivate fiscal policy to provide too much fiscal support ex ante. However, this effect is overlaid by other factors, which explains why fiscal policy is finally less pro-cyclical ex-post than fiscal plans. A possible explanation for the pro-cyclicality of fiscal plans could be for governments to signal a tighter policy stance than will actually be implemented in the end. The reasons for pro-cyclical fiscal policy, commenly documented in empirical studies, remains an area for further research.

Appendix

1.A Additional descriptive graphs and tables

Figures 1.A.1 and 1.A.2, and tables 1.A.1 and 1.A.2 replicate the figures and tables in section 1.4.1 with the change in the output gap instead of the level of output gap.

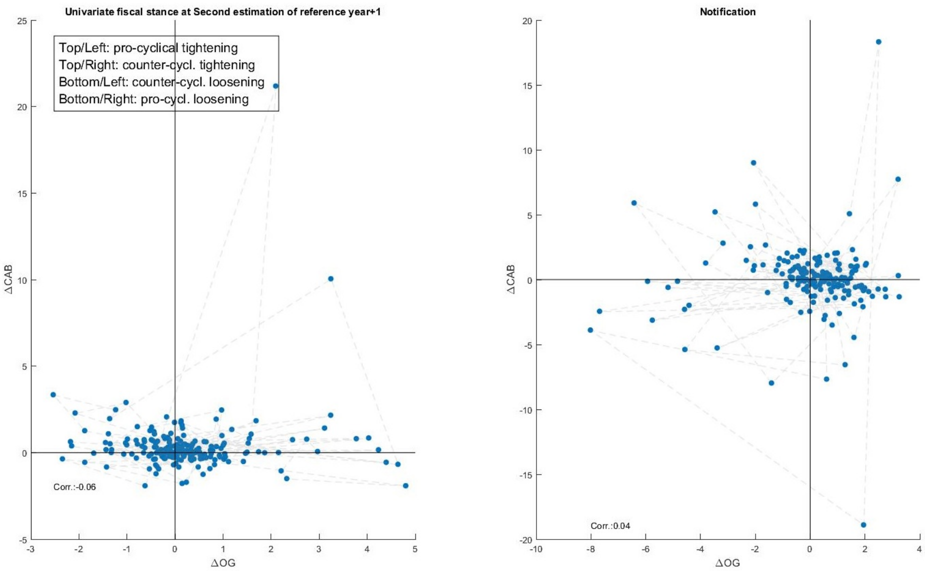


Figure 1.A.1: Fiscal stance throughout the cycle over 2004-2019 (ΔCAB and ΔOG)

Note: a tightening fiscal policy corresponds to an increase in the CAB

	Pro-cycl. tight.	Anti-cycl. tight.	Pro-cycl. loosen.	Anti-cycl. loosen.
$E_{t-1}(x_t)$	44	54	52	26
$x_{t\|t+1}$	38	53	61	24

Table 1.A.1: Fiscal stance regimes for ΔCAB and ΔOG over 2004-2019

	Pro-cycl. tight.	Anti-cycl. tight.	Pro-cycl. loosen.	Anti-cycl. loosen.
$E_{t-1}(x_t)$	47	47	59	23
$x_{t\|t+1}$	36	44	70	26

Table 1.A.2: Fiscal stance regimes for $\Delta CAPB$ and ΔOG over 2004-2019

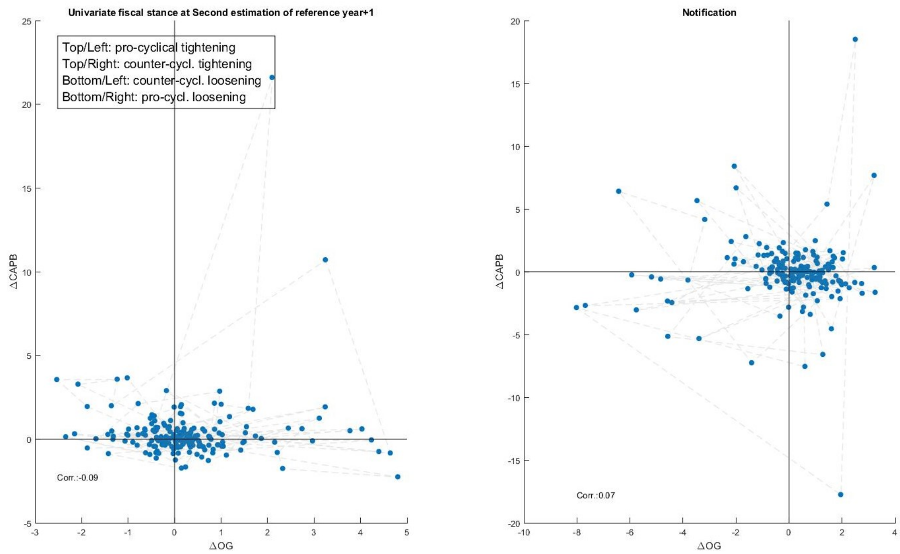


Figure 1.A.2: Fiscal stance throughout the cycle over 2004-2019 ($\Delta CAPB$ and ΔOG)

Note: a tightening fiscal policy corresponds to an increase in the CAPB

Figure 1.A.3 shows the distribution of output gap and cyclically-adjusted primary balance forecast errors over 2004-2019.

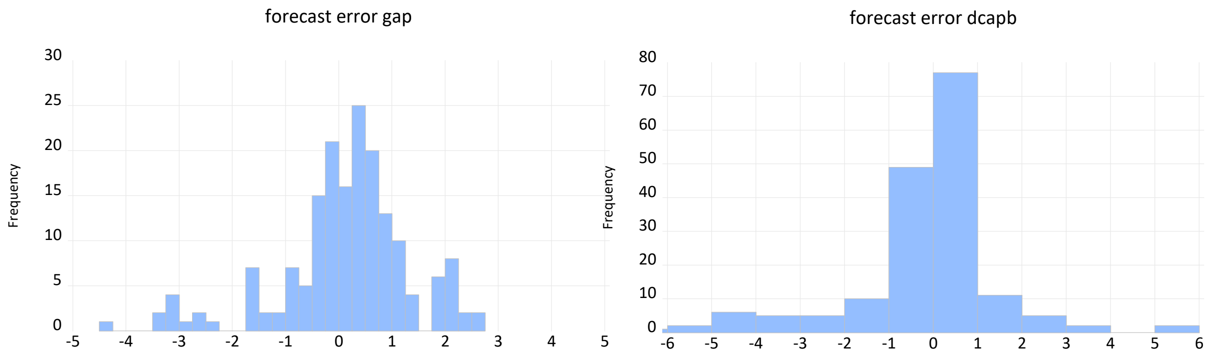


Figure 1.A.3: Distribution of output gap and cyclically-adjusted balance forecast errors, 2004-2019

Table 1.A.3 shows the forecast errors of ΔCAB and GAP grouped per country.

	AT	BE	DE	ES	FI	FR	GR	IE	IT	NL	PT
Error GAP	0.28	0.31	0.40	-0.42	0.01	0.31	-0.59	0.64	-0.30	0.27	0.45
Error dCAB	0.03	0.20	0.37	-0.72	0.13	-0.08	-0.61	-1.65	0.13	0.07	-0.03
Nb error GAP > 0	9	13	13	5	10	12	6	13	3	10	12
Nb error dCAB > 0	10	9	14	7	10	9	5	7	9	9	11

Table 1.A.3: Average forecast errors of ΔCAB and GAP

1.B Additional estimation results

	dependant: $\Delta CAB_{i,t}$		dependant: $\Delta CAPB_{i,t}$	
	(1)	(2)	(3)	(4)
Method	LS-FE	LS-FE	LS-FE	LS-FE
$CAB_{i,t-1}$ or $CAPB_{i,t-1}$	-0.332***	-0.382***	-0.398***	-0.441***
	[-2.650]	[-2.713]	[-3.014]	[-3.053]
$GAP_{i,t-1}$	-0.018	-0.038	-0.014	-0.003
	[-0.143]	[-0.214]	[-0.116]	[-0.018]
$DEBT_{i,t-1}$	0.036	0.033	0.036	0.046
	[1.561]	[1.124]	[1.560]	[1.541]
Constant	-3.860*	-3.807	-3.030	-3.849
	[-1.971]	[-1.545]	[-1.594]	[-1.582]
Nb. of obs.	176	176	176	176
Adj. R^2	0.180	0.267	0.208	0.309
Country FE	yes	yes	yes	yes
Time FE	no	yes	no	yes

Table 1.B.1: Fiscal reaction functions in the euro area (2004-2019) for $\Delta CAB_{i,t}$ and $\Delta CAPB_{i,t}$ with final data

Asterisks *, ** and *** denote significance at 10%, 5% and 1% significance levels, respectively. T-statistics are given in parentheses. Robust standard errors (cross-section SUR PCSE). Country and time fixed effects are not reported. The coefficient of $GAP_{i,t-1}$ is also not significant in (1)-(4) when using ordinary standard errors.

	dependant: $\Delta CAB_{i,t}$		dependant: $\Delta CAPB_{i,t}$	
Data vintage	Budget planning	First outcome	Budget planning	First outcome
Method	LS-FE	LS-FE	LS-FE	LS-FE
$CAB_{i,t-1}$ or $CAPB_{i,t-1}$	-0.474***	-0.352***	-0.541***	-0.406***
	[-6.668]	[-2.674]	[-8.467]	[-2.978]
$\Delta GAP_{i,t-1}$	-0.456**	0.183	-0.530**	0.211
	[-2.067]	[0.676]	[-2.610]	[0.797]
$DEBT_{i,t-1}$	0.023**	0.021	0.042***	0.030
	[2.100]	[0.871]	[3.929]	[1.191]
Constant	-2.564***	-2.357	-2.724***	-2.136
	[-2.752]	[-1.112]	[-3.059]	[-1.031]
Nb. of obs.	176	176	165	176
Adj. R^2	0.652	0.262	0.741	0.306
Country FE	yes	yes	yes	yes
Time FE	yes	yes	yes	yes

Table 1.B.2: Fiscal reaction functions in the euro area (2004-2019) for $\Delta CAB_{i,t}$ and $\Delta CAPB_{i,t}$ with the change in the output gap

Asterisks *, ** and *** denote significance at 10%, 5% and 1% significance levels, respectively. T-statistics are given in parentheses. Robust standard errors (cross-section SUR PCSE). Country and time fixed effects are not reported.

	Budget planning		First outcome	
	(1)	(2)	(3)	(4)
Method	LS-FE	LS-FE	LS-FE	LS-FE
$CAB_{i,t-1}$	-0.388***	-0.459***	-0.294***	-0.330***
	[-6.267]	[-7.200]	[-3.115]	[-2.906]
$GAP_{i,t-1}$	-0.046		0.123	
	[-0.627]		[1.164]	
$DEBT_{i,t-1}$	0.004	-0.000	0.017	0.023
	[0.441]	[-0.031]	[1.167]	[1.195]
$dummy_{GFC}$	0.159		-2.708***	
	[0.314]		[-4.045]	
$dummy_{EDC}$	-0.158		-0.138	
	[-0.321]		[-0.231]	
$GAP_{i,t-1} * dummy_{GFC}$		0.472		0.092
		[1.149]		[0.147]
$GAP_{i,t-1} * dummy_{EDC}$		-0.150		0.250
		[-1.505]		[1.130]
$GAP_{i,t-1} * (1 - dummy_{GFC} - dummy_{EDC})$		-0.187**		0.034
		[-2.226]		[0.208]
Constant	-0.751	-0.706	-1.371	-2.273
	[-1.030]	[-1.000]	[-1.159]	[-1.484]
Nb. of obs.	231	231	231	231
Adj. R^2	0.494	0.625	0.205	0.248
Country FE	yes	yes	yes	yes
Time FE	no	yes	no	yes

Table 1.B.3: Fiscal reaction functions in the euro area (1999-2019) for $\Delta CAB_{i,t}$ controlling for crisis years

Asterisks *, ** and *** denote significance at 10%, 5% and 1% significance levels, respectively. T-statistics are given in parentheses. Robust standard errors (cross-section SUR PCSE). Country and time fixed effects are not reported. The coefficient of $GAP_{i,t-1} * dummy_{GFC}$ and $GAP_{i,t-1} * dummy_{EDC}$ are weakly significant in (2) when using ordinary standard errors. They remain insiginificant in (4) in this case.

	Budget planning (1)	First outcome (2)
Method	LS-FE	LS-FE
$CAB_{i,t-1}$	-0.805***	-0.537**
	[-12.91]	[-2.32]
$GAP_{i,t-1}$	0.642***	1.002
	[3.83]	[1.43]
$DEBT_{i,t-1}$	-0.043**	0.025
	[-3.44]	[0.49]
$dummy_{PRO}$	-4.177***	-5.144
	[-4.61]	[-1.49]
$CAB_{i,t-1} * dummy_{PRO}$	0.446***	0.167
	[5.50]	[0.67]
$GAP_{i,t-1} * dummy_{PRO}$	-0.856***	-1.002
	[-4.75]	[-1.44]
$DEBT_{i,t-1} * dummy_{PRO}$	0.055***	0.036
	[4.21]	[0.67]
Constant	2.326	-0.684
	[2.95]	[-0.22]
Nb. of obs.	176	176
Adj. R^2	0.825	0.486
Country FE	interacted	interacted
Time FE	yes, interacted	yes, interacted

Table 1.B.4: Fiscal reaction functions in the euro area (2004-2019) for $\Delta CAB_{i,t}$ with pro-cyclical country group dummy

Asterisks *, ** and *** denote significance at 10%, 5% and 1% significance levels, respectively. T-statistics are given in parentheses. Robust standard errors (PCSE). Country and time fixed effects as well as their interaction effects are not reported. The dummy $dummy_{PRO}$ takes the value 1 for countries that have an ex ante pro-cyclical fiscal stance and 0 otherwise.

Table 1.B.5 provides the coefficients on the lagged output gap for individual fiscal reaction functions estimated for $\Delta CAPB_{i,t}$, as we did for $\Delta CAB_{i,t}$ in section 1.5.3. The conclusions are similar as those obtained for $\Delta CAB_{i,t}$ except that the coefficients for the Netherlands is now negative (although still non significant).

Group	Country	Budget planning	First outcome
Group 1 (- sign on $GAP_{i,t-1}$ in budget planning)	GR	-0.505**	0.097
		[-2.606]	[0.225]
	FR	-0.390**	-0.385
		[-2.911]	[-1.781]
	IT	-0.185	-0.042
		[-0.895]	[-0.281]
	ES	-0.161	-0.342
		[-1.665]	[-1.493]
	PT	-0.092	-0.057
		[-0.550]	[-0.210]
	NL	-0.043	-0.267
		[-0.206]	[-1.186]
	FI	-0.035	-0.121
		[-0.217]	[-1.606]
Group 2 (+ sign on $GAP_{i,t-1}$ in budget planning)	BE	0.010	-0.580**
		[0.072]	[-2.482]
	AT	0.297*	0.044
		[2.192]	[0.315]
	DE	0.587***	0.169
		[4.851]	[1.473]
	IE	0.920***	1.169
		[3.665]	[1.316]

Table 1.B.5: Individual fiscal reaction functions (2005-2019) for $\Delta CAPB_{i,t}$: coefficient on lagged output gap

Asterisks *, ** and *** denote significance at 10%, 5% and 1% significance levels, respectively. T-statistics are given in parentheses. Ordinary least squares. Ordinary standard errors.

The fiscal reaction functions estimated in panel of group-1 and group-2 countries for $\Delta CAPB_{i,t}$ are again similar to those obtained in section 1.5.3 for $\Delta CAB_{i,t}$: whatever the ex ante fiscal stance (procyclical in group 1, countercyclical in group 2), the fiscal policy is acyclical ex post.

Group of countries	Group 1 (pro-cyclical)		Group 2 (counter-cyclical)	
Data vintage	Budget planning	First outcome	Budget planning	First outcome
Method	LS-FE	LS-FE	LS-FE	LS-FE
$CAPB_{i,t-1}$	-0.441***	-0.459***	-0.837***	-0.553
	[-5.627]	[-3.461]	[-10.29]	[-1.614]
$GAP_{i,t-1}$	-0.182**	0.064	0.784***	1.156
	[-2.026]	[0.424]	[2.979]	[0.961]
$DEBT_{i,t-1}$	0.023**	0.063**	-0.008	0.063
	[2.247]	[2.573]	[-0.566]	[0.868]
Constant	-1.705*	-5.088**	1.974	-4.384
	[-1.961]	[-2.405]	[1.647]	[-0.790]
Nb. of obs.	105	105	60	60
Adj. R^2	0.613	0.421	0.885	0.292
Country FE	yes	yes	yes	yes
Time FE	yes	yes	yes	yes

Table 1.B.6: Fiscal reaction functions in subgroups of euro area countries (2005-2019) for $\Delta CAPB_{i,t}$

Asterisks *, ** and *** denote significance at 10%, 5% and 1% significance levels, respectively. T-statistics are given in parentheses. Robust standard errors (cross-section SUR PCSE). Country and time fixed effects are not reported.

Chapter 2

Fiscal Expenditure Spillovers in the Euro Area[1]

2.1 Introduction

Fiscal spillovers across countries have received increasing attention in recent years. Understanding the impact of a country's fiscal policies on macroeconomic conditions in other Member States in a monetary union, as the euro area, is of considerable interest to a central bank setting a single monetary policy. This allows the bank to better gauge the euro area's overall economic developments, and in turn feeds into the assessment of risks to price stability. Moreover, fiscal spillovers should be taken into account when assessing the aggregate euro area fiscal stance. Finally, the size of fiscal spillovers is important when assessing the stabilisation effects of national fiscal policies. If fiscal spillovers are small, then the existence of a central fiscal stabilisation function that can support national economic stabilisers in the presence of large economic shocks would make the Economic and Monetary Union (EMU) more resilient. In contrast, stronger spillovers would speak in favor of some coordination of fiscal policies among Member States, especially in the case of large symmetric shocks.

National fiscal policies spill over to other countries through different channels. Trade is an important transmission channel between countries, whereby a fiscal expansion in one country increases its imports from other countries. Fiscal expansion could also increase domestic prices and the real effective exchange rate, reinforcing spillovers, as the stimulating country loses competitiveness vis-à-vis the other countries (relative price channel). In a monetary union, the monetary policy reaction (interest-rate channel) depends on the impact of national fiscal action on aggregate price developments. Also, interest rates may

[1]This chapter is part of a broader project on spillovers using both empirical methods and DSGE modelling, with M. Alloza (Banco de España), M. Ferdinandusse (European Central Bank) and P. Jacquinot (European Central Bank). The joint work has been published as Occasional Paper of the European Central Bank in April, 2020. This chapter presents exclusively the work that I have conducted on my own as part of this project.

occasionally fail to react to price changes stemming from fiscal action, for instance, if the economy is constrained by the effective lower bound. The net spillover effect is hence ambiguous and requires empirical evidence to determine the spillover effect.

This chapter provides new empirical estimates of the magnitude of fiscal spillovers in the euro area, using a panel vector autoregression (PVAR) model based on annual data since 1972. The paper focuses on an expenditure-based (consumption and investment) fiscal policies. Expenditure-based and tax-based fiscal expansions show generally different domestic fiscal mulipliers. Due to the marginal prospensity to save, the Keynesian (short-term) multiplier is generally higher for spending increases than for net tax cuts. On the other hand, tax cuts are likely to have a positive impact on aggregate supply, and hence output in the long-term. We focus in this chapter on expenditure-based fiscal spillovers, which is due to data availability for fiscal variables for the euro area. While data for expenditure are readily available for the main euro area countries from the early 1970s, corresponding revenue data are scarse and only available since 1996, which greatly reduces the robustness of the estimates. We also compare the fiscal spillovers in size to domestic multiplier to gauge their relative importance to domestic fiscal policy action.

Throughout the chapter, we employ the concept of destination spillovers, compared to spillovers by origin. Destination spillovers measure spillovers from a spending shock in other countries on a specific recipient country. This indicates the magnitude of spillovers from common fiscal action among euro area Member States to another Member Sate. In contrast, spillover by origin would show the impact on the output in the receiving countries of fiscal action in one specific Member State.

Our empirical findings underline that fiscal expenditure has positive and non-negligible spillover effects across countries in the euro area. The magnitude of fiscal spillover lies however far below those of domestic fiscal multipliers and depends upon the size and trade openess of countries where the fiscal shock originates. We find that fiscal spillovers are larger for more open and closely integrated economies and smaller if generated in relatively larger economies. There exist also important state dependencies in fiscal spillovers that stem from the cyclical position of the economy, the reaction of monetary policy and/or other factors. This state dependency makes it difficult to take spillover effects effectively into account in the conduct of domestic fiscal policy.

The remainder of the paper is organised as follows: Section 2.2 reviews the existing literature and highlights our contribution to it. Section 2.3 describes the employed data and section 2.4 our empirical methodology. In section 2.5 we present and discuss the results for the benchmark estimation and a number of variantes for different country and time subsamples. In addition, we present several robustness checks. The final section 2.6 concludes.

2.2 Literature review

While the empirical literature on fiscal spillovers has grown in recent years, it remains relatively limited. Differing identification of fiscal shocks and presentation of the results according to different metrics makes generalising the findings from the literature all the more complex. In addition, there are few papers that focus exclusively on the euro area. Our paper aims to bridge this gap and provide systematic empirical spillover estimates, comparable to standard domestic Keynesian fiscal multipliers, for the euro area.

In the earliest work closest to our approach, Beetsma et al. (2006) and Beetsma and Giuliodori (2011) have analysed fiscal spillovers from government spending shocks in 14 countries of the European Union using a Panel VAR approach. For example, based on annual data from 1965 to 2004, Beetsma et al. (2006) estimate that a spending-based fiscal expansion of 1% of GDP in Germany would lead to an average increase in the output of other EU economies by 0.15% after two years; for an expansion originating in France, the impact is 0.08%. Beetsma and Giuliodori (2011) concentrate mainly on the consequences of fiscal policy on domestic activity over the period from 1970 to 2004, employing also a Panel VAR approach. In this framework, they show also show however that a government spending shock originating in the five largest EU economies (France, Germany, Italy, Spain and the UK) increases output in the rest-of-EU by up to 0.35 (peak effect) after two years.

Auerbach and Gorodnichenko (2013) provide cross-country spillover effects of government purchases[2] for output in a large number of OECD countries. They identify the fiscal shock as one-period-ahead forecast error for government purchases from fiscal reaction functions (normalised to the recipient country's GDP) and employ a local projection (LP) method for different estimation periods. In their baseline result, they document large spillovers effects of government purchases (1.34 using time-varying weights and 1.60 using fixes-weights for the average three-year effect). A comparison of the estimated spillover effects with domestic multipliers using the same identification and estimation strategy is unfortunately not available in the paper. Auerbach and Gorodnichenko (2013) also show that the magnitude of the spillover varies with the state of the economy in the recipient and source countries, with the multipliers being larger in recessions and modest in expansions. Goujard (2017) also estimates fiscal spillovers for consolidation episodes across OECD countries over the 1978 to 2009 period, employing the LP method and a narrative approach of shock identification. He finds that a fiscal contraction in Germany by 1% of GDP is associated with a reduction of output by 0.23 % for a typical OECD country (spillover by origin).

The International Monetary Fund (IMF) has provided more recent empirical estimates of fiscal spillovers (Blagrave et al., 2017; Dabla-Norris et al., 2017; IMF, 2017; Poghosyan, 2017). Using quarterly data from 2000 to 2016 for 55 countries, IMF (2017) and Blagrave

[2] Government purchases are the sum of government consumption and government investment, which is equivalent to government spending minus social transfers, see also chapter 2.3.

et al. (2017) report that a one percent of GDP fiscal stimulus in a major advanced economy (France, Germany, Japan, the US and the UK) can raise output in recipient countries by 0.08% over the first year both for government and tax shocks. Spillovers for government spending shocks are larger: output in recipient countries can increase by 0.15% following a spending hike, against 0.05% after a tax cut. They also show that spillovers are larger when: cyclical conditions are weak; monetary policy does not counteract fiscal shocks (for example, when monetary policy is constraint by the effective lower bound); when exchange-rate adjustments are limited (fixed-exchange rate regimes our currency unions) between source and recipient countries. When restricting the sample to Europe, the IMF estimates the one-year spillover of a 1% government spending shock in Germany and in France to be 0.26% and 0.14%, respectively (Blagrave et al., 2017). Dabla-Norris et al. (2017) estimate spillover effects for 10 euro area countries using quarterly data over the period from 1999 to 2016, based on a Panel VAR method with sign restrictions. They find positive cross-country spillovers, that can be large and significant. Larger economies and countries that are more integrated generate more sizeable spillovers. For example, a fiscal shock in Germany expands output in France at year one by 0.09, and by 0.17 in the Netherlands and 0.16 in Austria. Similarly, Poghosyan (2017) assesses spillovers from fiscal consolidations in 10 euro area countries from 1980 to 2015 using LP incorporating a spatial lag term. This approach allows to gauge the importance of direct and indirect spillover effects separately. On average, a 1% fiscal consolidation reduces total output by 0.6% on impact, out of which 0.32% is explained by fiscal spillovers.

Alloza et al. (2019) provide individual estimates of government spending spillovers for the four largest euro area countries based on quarterly data for the period 1980-2016 using LP. They find an average destination spillover effect (cumulative multiplier) after two years of 0.73 for Germany, 0.50 for France, 0.32 for Spain and 0 for Italy (for a shock originating in the respective other three countries) and an average output-weighted spillover effect over the four euro area countries of 0.46. Alloza et al. (2019) show that there is a substantial and positive relationship between the size of cross-country government spending spillovers and their corresponding share of bilateral imports.

Overall, the existing empirical literature suggests that fiscal spillovers are positive, while the size of spillovers varies depending on the methodology employed and countries under consideration and a number of other factors.

Additional insight into fiscal spillovers and interactions is provided by a number of studies using other methods, such as general equilibrium models (see for example in't Veld, 2013 and in't Veld, 2017) and spatial econometric models. Rich DSGE multi-country models can provide more insight into the determinants of fiscal spillovers than empirical methods such as VAR, which encompass a variety of contributing effects that are difficult to disentangle. However, DSGE models may come at the price of imposing restrictive assumptions, which may not always have strong empirical foundations. Studies based on DSGE models often find spillovers in normal times to be lower than the VAR-based estimates, but higher when interest rates do not react. in't Veld (2013) finds that a (temporary two year) 1 percent of GDP increase in government investment in Germany and the rest of the euro area core increases the real GDP in other euro area countries by between 0.2% and 0.3%. Barbier-Gauchard and Betti (2020), based on a two-country DSGE model of monetary union with an extendend labour market bloc, find that spillovers effects differ widely according to the fiscal instrument. For the expenditure categories, a rise in transfers to households and in public investment produce positive spillover effects on foreign output, while a rise in public consumption generates negative spillover effects.

Using a slightly different angle of fiscal interactions between OECD countries, Cassette et al. (2013) show with the help of spatial econometric methods that discretionary fiscal policies implemented in one country not only influence output in other countries but also discretionary fiscal actions in other countries. The effects are stronger the closer countries are located geographically and the stronger they are linked economically.

2.3 Data

The panel estimates are based on an annual panel dataset of 11 euro area countries. The countries covered are the initial euro area countries excluding for Luxembourg[3], from 1972 to 2017. The benchmark model includes five variables: trade-weighted foreign real government spending (govfor), i.e. the sum of real government consumption and real government investment[4], domestic real government spending (gov), real GDP (gdp), the GDP deflator (def) and the long-term (10-year) nominal interest rate (ltn). An alternative specification of the model adds the real effective exchange rate (reer) to the model as a control for the reaction of relative prices. All variables are in logs (apart from the nominal interest rate) and in 1st difference. The variables and data sources are listed in table 2.1.

Foreign government spending. To calculate aggregate foreign real government spending, real government spending in euro area partner countries is weighted by bilateral import weights. Although a large part of government spending consists of non-traded goods, weighting government spending by trade flows provides a synthetic indicator that reflects the relative exposure of the domestic economy to fiscal shocks in other euro area countries, which is standard in the literature (see Beetsma et al., 2006). We adopt a destination country perspective, i.e. we calculate by how much trade-weighted real government spending in euro area source countries affects the destination country's output, prices, interest rates etc. More precisely, the foreign spending variable is calculated as

[3] Austria, Belgium, Germany, Spain, Finland, France, Greece, Ireland, Italy, Netherlands and Portugal. Data limitations do not allow for the inclusion of Luxembourg and later euro area entrants

[4] A chain-type Fisher index is used to construct real government spending from government investment and government consumption, using the part of each component in total nominal government spending over the years t-1 and t as weight.

follows:

$$govfor_{i,t} = \sum_{j=1}^{J=10} \frac{M_{ij,t-1}}{M_{j,t-1}} gov_{j,t} \tag{2.1}$$

where i denotes the recipient country, j the source country, $M_{ij,t-1}$ country j's good imports from i at time t-1 (10-year moving average), $M_{j,t-1}$ country j's total good imports at t-1 (10-year moving average), and $gov_{j,t}$ government spending in source country j. Import flows are averaged over ten year to avoid endogeneity problems and avoid idiosyncratic shocks to influence the results.

Trade weights sum up to 1 for each country j over i ($i \neq j$), i.e. we presume that euro area countries only trade with each other (government shocks are not exported to non-euro area countries). This assumption is due to the limited availability of bilateral trade data over a long horizon with non-euro area countries.[5] The estimated spillover effects are hence likely to represent upper bounds, as a part of the spending shocks should also be exported to non-euro area countries. Figure 2.1 shows foreign real government spending fluctuations for the average euro area country for the period from 1972 to 2017. They were higher at the beginning of the period up to the 1990s (average fluctuations of 5.2%) before falling to around 1.9% on average for the period after 1990. A fall in the average foreign (trade-weighted) real government spending is observed following the Great Recession.

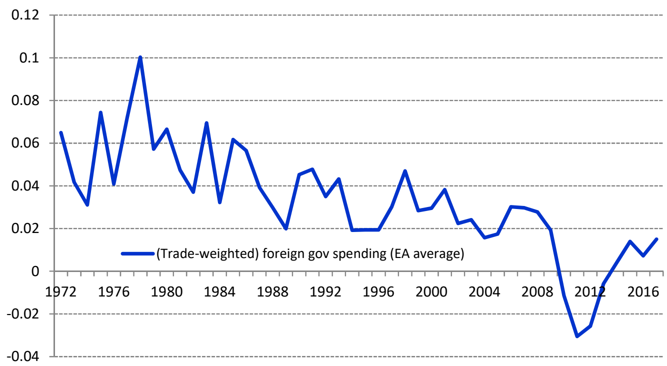


Figure 2.1: Trade-weighted foreign government spending (euro area average)

[5]We accept this assumption in this context as we are interested in intra-euro area spillovers.

Variable	Source	Starting date	Observation
Real government spending			Sum of government consumption and investment, obtained by a Fisher chain
Real government consumption	Ameco	1971	
Real government investment	Ameco	1971	Up to 1994, deflated by the total investment deflator, from 1995 government investment deflator
Real GDP	Ameco	1971	
GDP deflator	Ameco	1971	
Real effective exchange rate	Ameco	1971	Against former EU-15, based on unit labour costs for the total economy
Nominal 10-year interest rate	IMF, Eurostat	1971	Eurostat when available, otherwise IMF. For countries where the data series starts after 1971 (ES, FI, GR, IT), the average spread of the first three available years is applied to the euro average rate
Potential GDP	Ameco	1971	
Bilateral imports	Eurostat	1988	10-year moving averages

Table 2.1: Data sources

2.4 Methodology

We estimate a Panel Vector Autoregression (PVAR). The Panel VAR approach combines the advantages of a VAR model, i.e. the ability to capture the intertemporal dynamics of the data imposing only a minimum of restrictions, with the advantages of a cross-country structure. This makes it a good tool for the analysis of the transmission of government spending shocks across countries and across time. Cross-country interdependency is captured in model via trade links and the unobserved individual country heterogeneity using fixed effects.

In the benchmark case, the model is estimated from 1972 to 2017. We also estimate the model over two shorter time spans from 1972 up to the Great Recession in 2007, in order to analyse whether benchmark results are influenced by the simultaneous fiscal action in many euro area countries during the Great Recession, and starting later in the 1980s (from 1980 to 2017). We estimate the PVAR for the 11 euro area countries and for the four largest euro area countries only.

Model. The main structure of the PVAR follows the general form as shown in Canova and Ciccarelli (2013):

$$Y_{i,t} = A_{0,i} + A_i(L)Y_{i,t-1} + u_{i,t} \tag{2.2}$$

where $Y_{i,t}$ is the vector of endogenous variables $Y_{i,t} = (govfor_{i,t}, gov_{i,t}, gdp_{i,t}, def_{i,t}, ltn_{i,t})'$ in the benchmark specification, $A_i(L)$ is a polynomial in the lag operator L, $A_{0,i}$ is a vector of country-fixed effects, and $u_{i,t}$ are the identically and independently distributed errors. $i = 1, ..., 11$ and $t = 1, ..., T$.

We estimate a parsimonious model with five variables in the benchmark specification, including foreign and domestic real government spending, real GDP, the GDP deflator and a nominal interest rate (the 10-year government bond rate). To check for robustness, we also estimate an alternative specification which includes the real effective exchange rate to control for the endogenous reaction of relative prices. The variables enter the model in first differences in order to satisfy the stationarity condition (no root lies outside the unit circle). The benchmark specification is estimated with 2 lags, as selected by the Hannan-Quinn information criterion and by t-statistics for residual serial correlation, and an alternative specification with 1 lag, as selected by the Bayesian information criterion.

Estimation. The model is estimated using a least square dummy variable (LSDV) estimator. Country-specific (time-invariant) fixed effects, which capture invariant unobserved country heterogeneity, are obtained by adding dummy variables to the system. The existence of lagged dependent variables in the model introduces a bias into the fixed effects estimator, whose size is however decreasing in terms of the number of time-series observations T (see Bun and Kiviet, 2006 and Nickell, 1981). This bias should be here, as T is large for a panel model (46 years) compared to the number of units N (11 countries).[6].

Identification. A recursive identification scheme through a Cholesky decomposition is used to identify structural shocks (see Blanchard and Perotti, 2002). The recursive form determines which variable responds to the structural shocks in time (t). In the benchmark specification, foreign government spending is ordered first, affecting all the other variables contemporaneously. Domestic spending, ordered second, is allowed to a have a contemporaneous impact on domestic GDP, prices and interest rates, but would affect foreign government spending only with a lag. This ordering is consistent with other VAR studies and seems plausible in the situation of one-year ahead annual budget planning, which leads to a certain inertia in government choices as to macroeconomic conditions (see Beetsma and Giuliodori, 2011). Nonetheless, we test for alternative orderings of the variables, ordering notably foreign government spending - which is our main variable of interest - last. Error bands of the impulse responses are generated by Monte Carlo simulations over 200 draws with confidence bands at 68%. The impulse response functions (IRFs) are reported for a one-standard deviation shock for a period of over four years.

[6] Another option would be to use the forward mean-differencing transformation - also known as Helmert procedure - to remove fixed effects and estimate the coeffcients by the GMM estimator. However, the GMM estimator is found to be only consistent for panels with a short time dimension, i.e. $N > T$, which is not the case here.

Multipliers. The estimated IRFs do not directly show multipliers, as the model is estimated in logs. We use the transformation initial proposed in Blanchard and Perotti (2002), and described more in detail in Ramey and Zubairy (2018), to transform the impulse responses into multipliers. As multiplier we understand the cumulative effect of a €1 government shock on the macroeconomic variable of interest. For this, we first transform the impulse responses based on the standardised shocks into elasticities (or one-unit impulse responses) by dividing them by the standard deviation of the shocks in the reduced form (including its dynamic responses). Secondly, we multiply these elasticities by the sample average GDP-to-government spending-ratio to obtain one-unit multipliers. The average ratio of GDP/GOV is 5.6 over the full sample period 1972-2017. As Ramey and Zubairy (2018) and others highlight, multipliers are sensitive to the calculation method. We rely on this conversion method, as it is straightforward to implement in the PVAR framework and allows for a comparison with papers closest to ours, that rely on the same method.

Fiscal Foresight. A general problem with the estimation of macroeconomic effects of government spending shocks using the above method is the so-called fiscal foresight problem (see Leeper et al., 2013). This problem arises when economic agents do not react to actual government action but rather to the annoucement of future action. In this case, the identified structural shocks would not represent unanticipated changes in government spending. There are different methods to overcome the fiscal foresight problem, for example the inclusion of news variables or government spending forecasts into the model (see Auerbach and Gorodnichenko, 2012). We believe however that the fiscal foresight problem is of lesser relevance in this case for two reasons: (*i*) we rely on annual data which reduces the possibility to anticipate government compared for shorter time horizons, and (*ii*) the focus of this paper lies on spillover effects from government action in other countries, which should be more difficult to anticipate by the agents in the destination country.

2.5 Empirical results

2.5.1 Main empirical results

The estimates provide evidence of a positive average destination spillover effect in the euro area across all countries in the panel. Figure 2.2 plots the (non-cumulative) output response in an average euro area country for a one-standard deviation (equivalent to 2.6%) increase in foreign (trade-weighted) real government spending in the other ten euro area countries. The solid lines gives the point estimate and the dashed lines indicate the 68% confidence bands. Output in the destination country increases by (non-cumulative) 0.3 percentage point in the first year and by 0.1 percentage point in the second year after the shock (effects are significant at the 68%-level), before decreasing gradually to zero thereafter. Government spending in the destination country does not react significantly to the foreign government spending shock, which assures that the reaction of domestic GDP is not driven by a simultaneous impulse from domestic government spending. Prices in the destination country also react positively to a foreign (trade-weighted) real government spending shock, showing a significant impact of 0.4 percentage point starting from the second year and remaining at this level in the third and fourth year. Long-term interest rates in the destination country show a significant negative reaction on impact, which turns positive in the second year.

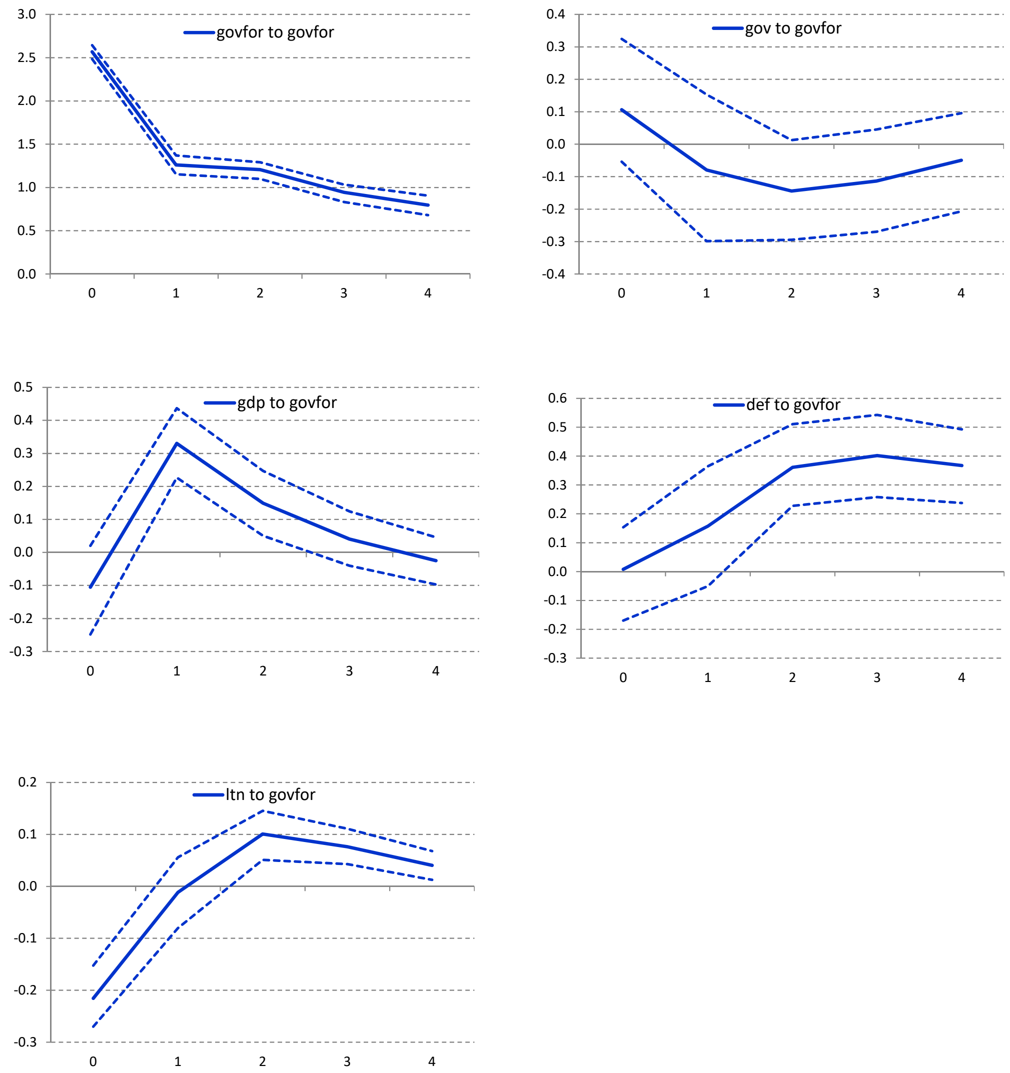


Figure 2.2: Non-cumulative IRFs for trade-weighted foreign government spending

Note: x-axis: years, y-axis: percentage points; average (non-cumulative) response in a euro area country to a trade-weighted increase in real government spending in the other euro area countries, 5-variables PVAR, 2 lags, LSDV, 68% confidence bands from Monte-Carlo simulations, 1972 to 2017.

For comparison, the domestic output response to a domestic government spending shock is 0.5 percentage point on impact, 0.3 percentage point in the first and second year and 0.1 percentage point in the third year, all significant at the 68%-level (see figure 2.3). Prices react, contrary to what could be expected, negative on impact and in the first year, before showing the expected positive sign starting from the second year. The reaction of long-term interest rate shows the well-know J-pattern[7].

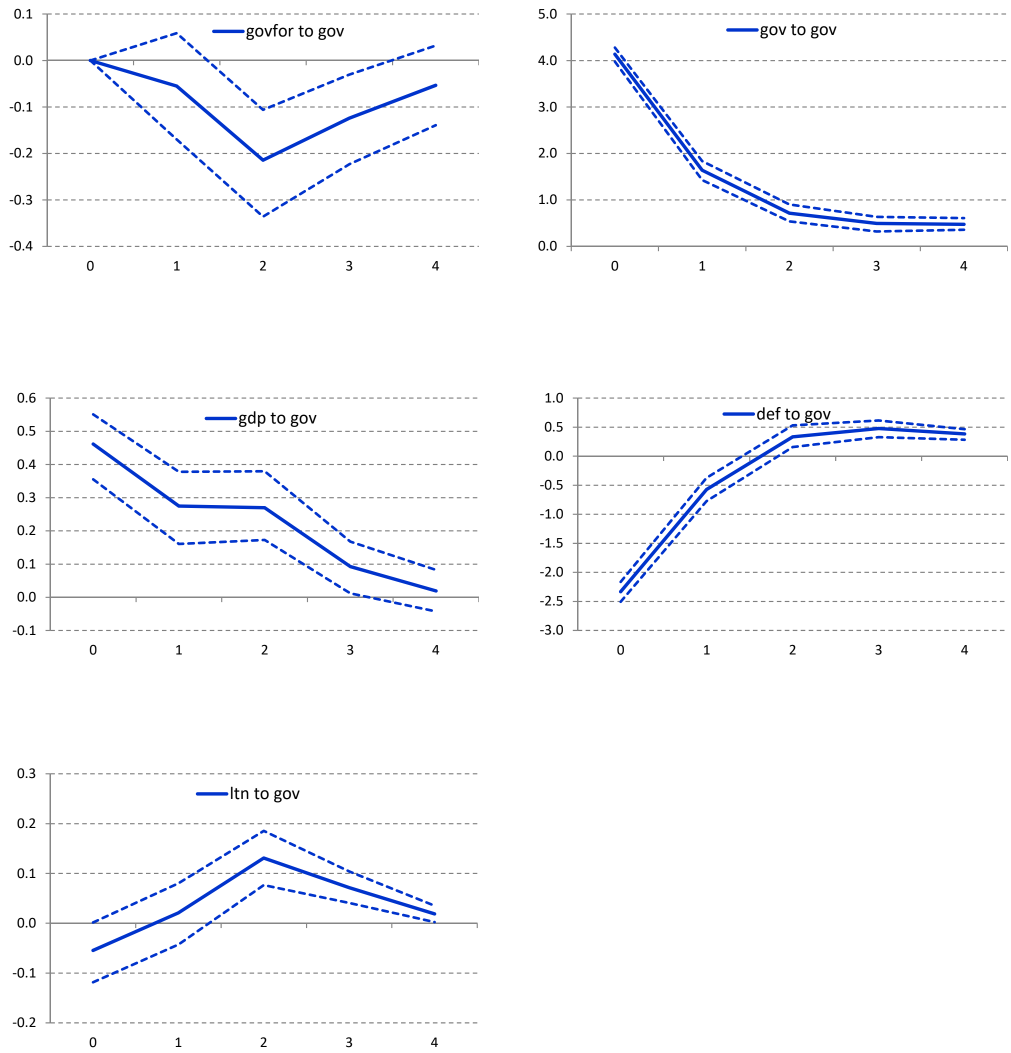


Figure 2.3: Non-cumulative IRFs for domestic government spending

Note: x-axis: years, y-axis: percentage points; average (non-cumulative) response in a euro area country to an increase in domestic real government spending, 5-variables PVAR, 2 lags, LSDV, 68% confidence bands from Monte-Carlo simulations, 1972 to 2017.

Cumulating these estimated output effects over time and dividing them by the cumulative impulse of government spending in the stimulating countries transforms these estimates into cumulative elasticities. We transform the elasticities into euro multipliers

[7]The initial fall of long-term interest rates could be explained by the positive output effect from government spending which dominates the worsened budget effect, while on the medium-term budget sustainability considerations could gain in importance

by multiplying them with the sample average GDP-to-government spending-ratio, as described in section 2.4. For a €1 foreign government spending shock, we find a cumulative destination spillover multiplier of €0.33 after one year, €0.42 after two years (the latter effect being significant at the 68%-level), decreasing to €0.32 after four years (see left-hand panel in figure 2.4. The cumulative domestic spending multiplier is €0.62 on impact, €0.71 after one year and €0.87 after two years, before decreasing slightly to €0.84 after four years, all effects being significant at the 68%-level (see right-hand panel in figure 2.4). These results imply domestic spending multipliers just below unity, and across countries of around 0.4.

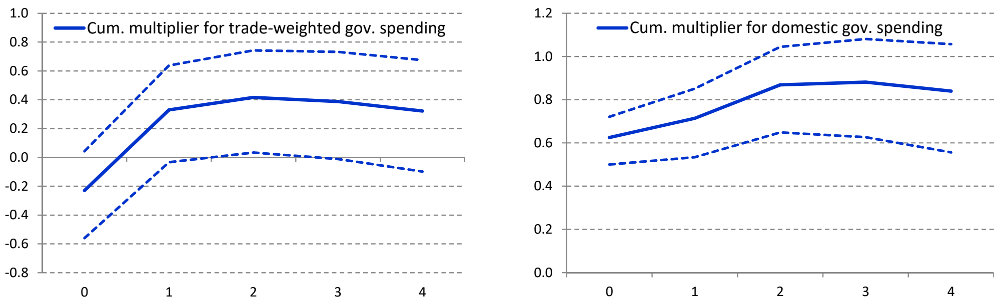


Figure 2.4: Cumulative multipliers for trade-weighted foreign and domestic government spending

Note: x-axis: years, y-axis: euro;, average cumulative multipliers in one euro area country of a trade-weighted €1 increase in government spending in the other euro area countries (left) and to a €1 increase in domestic government spending (right).

Our results lie within the range of previously estimated spillover effects in Europe, even if results are not always fully comparable due to different calculation methods for the multipliers and different country samples. Our estimated multiplier is slightly higher than the one reported in Beetsma et al. (2006). They find a spillover effect of 0.15 for a 1% of GDP spending shock originating in Germany on the output of other EU economies after two years and of 0.08 for a spending shock originating in France. Our slightly higher spillover estimate may be due to the fact that our panel also includes some small, very open economies, which can be expected to generate higher spillovers. We look in more detail into this question in the next section. Our estimate ist very close to Beetsma and Giuliodori (2011), who find a peak spillover effect of 0.35 after two years for a spending shock originating in the largest 5 EU economies on the output of the rest of the EU. Dabla-Norris et al. (2017) show bilateral spillover effects for 10 euro area countries which lie between 0.01 and 0.49 after one year. At the upper end of the range of the empirical estimates, Alloza et al. (2019) find a trade-weighted average destination spillover effect of 0.46 after two years for the four largest euro area economies, increasing to 0.60 after the third year, which is higher than our estimate.

2.5.2 Country characteristics and estimation period

Several papers find that the magnitude of fiscal spillovers depends upon the size of the country where the fiscal shock originates and the commercial and financial links between source and destination country. Dabla-Norris et al. (2017) find larger spillover for larger source countries and for countries that are more closely integrated. Beetsma et al. (2006) find a larger spillover effect from a government spending shock originating in Germany for small open economies such as Austria, Belgium and the Netherlands sharing a land border with Germany. Compared to the average spillover effect of 0.15 mentioned above, the spillover effect in this case increases to 0.4 after two years. To examine this question, we split our panel into two subsamples, one with the four large euro area countries (Germany, France, Italy and Spain) and one with the 6 smaller and generally more open economies[8] The four largest euro area countries have an average import share (imports from the euro area) of 10% of GDP over the 1972-2017 period against 26% of GDP for the smaller countries in the panel. The small country sample includes in particular very open economies such as Belgium, Austria and Netherlands. Other factors than trade openess such as differences in the composition of government spending across countries[9] or a stronger response of interest rates to fiscal shocks in large countries might also explain some of the differences.

We find that spillovers are lower when limiting the panel to the four largest euro area countries. For a trade-weighted one-standard deviation government spending shock in the three largest countries, output in the remaining country increases on average by only 0.1 percentage point in the first year (not significant at the 68% level) and by 0.2 percentage point in the second year (significant at the 68% level) after the shock (see left-hand panel of figure 2.5). However, owing to a stronger negative output response on impact, cumulative multipliers are close to zero after two years, increasing only to €0.15 after four years (against 0.32 in the full panel), see table 2.2. The estimate is less precise compared to the full sample and none of the multiplier effects is significant at the 68%-level.

In a next step, we rerun the exercice using a sample which includes the four large economies as source countries but all euro area countries in the destination sample in order to examine whether the result is driven by the size of the source countries (as compared to the size of the desitination countries). The effect is comparable in size to the one obtained with the sample including only the four large economies in the source and destination sample. It shows a cumulative euro spillover multiplier of €0.16 after four years, confirming the hypothesis that the size of spillovers depends on the size and trade openness of the source countries.

[8] Compared to the full sample, we exclude Greece from the split sample as Greece has a much lower trade openess than the other small countries.

[9] Both country groups show a similar composition of government investment vs. government consumption in total government spending in the full sample, but differences might exist on a more disaggregated level btween the two country groups

In the sample with the small countries, we find a positive and significant spillover effect on ouput of 0.3 percentage point in the first year and 0.2 percentage point in the second year, before turning negative in the third year (see right-hand panel in figure 2.5). The cumulative spillover multiplier is high, close to €1 after two years and €0.60 after four years (against 0.32 in the full panel), see table 2.2. We conclude that small, open economies generate larger government spending spillovers, while the less open larger economies produce smaller spillovers, and this independantly of the size of the destination countries. While the finding that more open economies generate higher spillovers confirms the finding of the literature, our results do not support the finding by Dabla-Norris et al. (2017) that fiscal shocks generating in larger countries create larger spillovers.

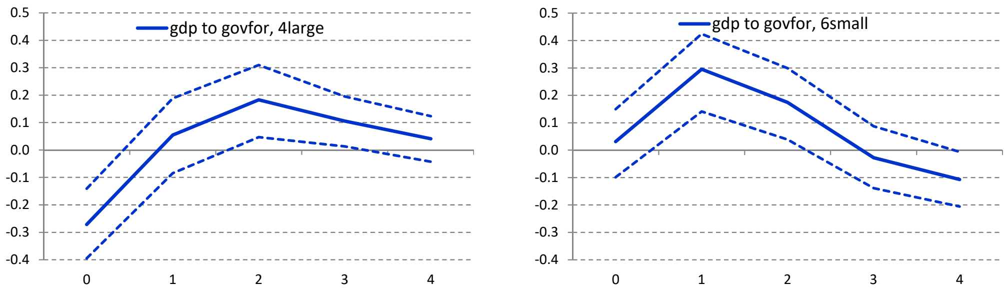


Figure 2.5: Non-cumulative IRFs for trade-weighted foreign government spending, large and small economies

Note: x-axis: years, y-axis: percentage points; average (non-cumulative) response in a euro area country to a trade-weighted increase in government spending in the other euro area countries, for four largest economies (left) and six small, open economies (right), 5-variables PVAR, 2 lags, LSDV, 68% confidence bands from Monte-Carlo simulations, 1972 to 2017.

In what regards domestic multipliers, one would expect them to be larger in the four larger countries, provided that less of the fiscal stimulus is exported to other countries. On the other hand, monetary policy is more likely to react to fiscal action in the larger countries[10], thus counterbalancing potentially the government spending shock and leading to smaller domestic multipliers in larger countries. Our results are not fully conclusive on this question. Domestic multipliers are slightly smaller for smaller economies in our sample compared to the larger ones. The average cumulative domestic multiplier for smaller economies is €0.43 on impact (vs. €0.62 in the full sample), €0.71 after two years (vs. €0.87 in the full sample) and €0.71 after four year (vs. €0.84 in the full country sample). All responses are significant at the 68%-significance level. However, we also find smaller domestic multipliers for the larger economies, where the cumulative domestic multiplier is €0.31 on impact, €0.62 after two years and €0.55 after four years

[10] This concerns both the the period since the start of the European monetary union in 1999 but also the period before 1999 as smaller economies had more years with pegged exchange rates

(all significant at the 68%-level). The slightly lower domestic multipliers in the two country subsamples compared to the total sample can be explained by the exclusion of Greece from the subsamples.

Going further, we carry out the estimations over two subperiods: (*i*) the period before the Great Recession from 1972 to 2007, and (*ii*) the period after the creation of the European Monetary System (EMS) from 1980 to 2017[11]. The first subsample analyses whether the synchronised fiscal action during the Great Recession affects the spillover results. This subsample also excludes the period of negative average real spending growth (see 2.1) during the Euro Area Sovereign Debt Crisis from the analysis. The second subsample compares the benchmark results to a period which excludes the 1970s, a time when European currencies were less closely aligned.

The results for the different subperiods confirm the positive destination spillovers. Figure 2.6 (left-hand panel) shows the (non-cumulative) output response to a trade-weighted foreign government spending shock in the estimation for the shorter sample from 1972 to 2007. The output response is comparable to the full sample in the first year, but it tends towards zero much faster thereafter (impact close to zero from the second year). Cumulative multipliers[12] are therefore only slightly smaller than in the full sample after two years (0.25 against 0.42) but significantly smaller after four years (0.02 against 0.32). These differences might be related to state dependencies in fiscal spillovers that stem from the cyclical position of the economy, the reaction of monetary policy and/or the negative fiscal shocks that are only observed in the years during and after the Great Financial crisis. Excluding notably the Great Financial crisis from the sample, seems to lead to smaller spillover effects.

In the second subsample from 1980 to 2017, the output effect of a foreign government spending shock is slightly smaller compared to the full sample in the first year, but it is more persistent with a significant and positive effect in the second and third year (see right-hand panel of figure 2.6). In cumulative multiplier terms, the effect is close to the full sample after two years (0.38 against 0.42 in the full sample) and slightly larger after four years (0.43 against 0.32 in the full sample), see table 2.2.

2.5.3 Fiscal instrument

In this section, we examine the size of spillovers separately for government consumption and government investment. Barbier-Gauchard and Betti (2020) find that spillovers differ widely according to the fiscal instrument. The spillover effect should depend significantly on the import content of the different fiscal goods and services.

We find indeed a slightly higher spillover multiplier for government consumption and a slightly lower spillover multiplier for government investment compared to the total spillover multiplier for the full sample from 1972 to 2017. For a trade-weighted one-

[11] The subsample periods were also selected to retain a sufficiently large number of degrees of freedom

[12] Multipliers are calculated using the average GDP-to-government-ratio over the reduced sample period

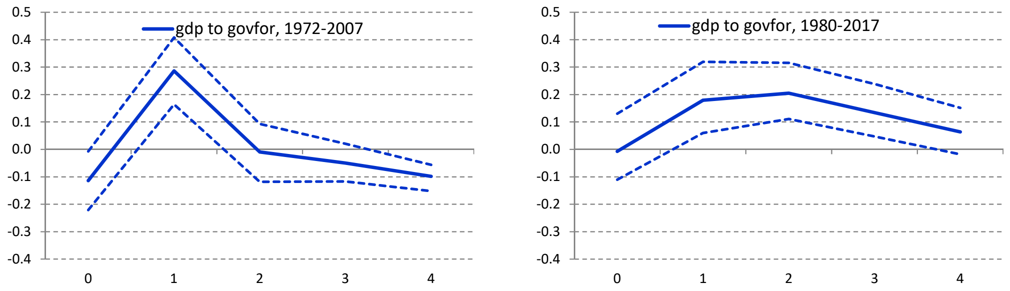


Figure 2.6: Non-cumulative IRFs for trade-weighted foreign government spending, different sample periods

Note: x-axis: years, y-axis: percentage points; average (non-cumulative) response in a euro area country to a trade-weighted increase in government spending in the other euro area countries, 5-variables PVAR, 2 lags, LSDV, 68% confidence bands from Monte-Carlo simulations, 1972 to 2007 (left) and 1980 to 2017 (right).

period	benchmark	4 Large	6 Small	1972-2007	1980-2017
0	-0.23	-0.87*	0.13	-0.27*	-0.01
1	0.33	-0.50	0.87*	0.30	0.22
2	0.42*	-0.05	1.04*	0.25	0.38*
3	0.39	0.11	0.85	0.16	0.43*
4	0.32	0.15	0.59	0.02	0.43*

Table 2.2: Cumulative multipliers for a €1 increase in trade-weighted foreign government spending

Note: *indicates significance at the 68% significance level

standard deviation government consumption shock, output in the destination country increases by 0.1 percentage point on impact, (non-cumulative) 0.1 percentage point in the first year and again 0.1 percentage point in the second year before decreasing gradually to zero until the fourth year (only the second year effect is significant at the 68% level), see left-hand panel of figure 2.7. The cumulative multiplier is €0.51 after four years, somewhat higher than the €0.32 multiplier after four years for total government purchases. Concerning government investment, the spillover effect is stronger concentrated in a specific year, with a negative foreign ouput effect on impact (-0.1 percentage point), a positive (non-cumulative) 0.25 percentage point increase after one year (the only year where the effect is significant at the 68% level), a 0.1 percentage point increase after two years, before falling quickly to zero afterwards, see right-hand panel of figure 2.7. The cumulative multiplier is €0.14 after four years.

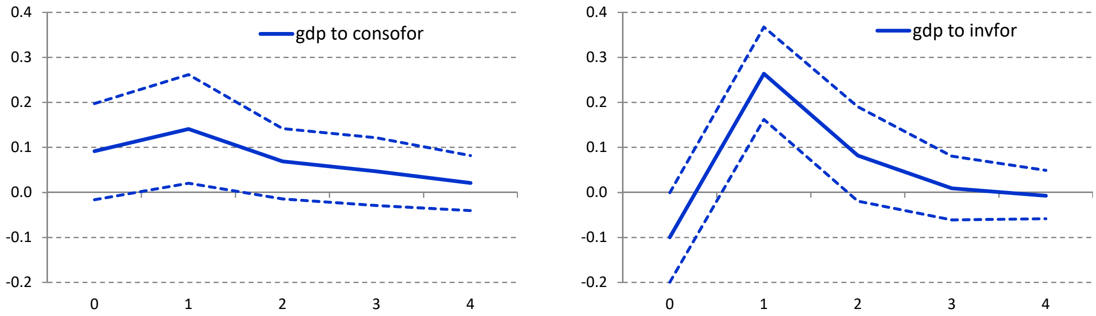


Figure 2.7: Non-cumulative IRFs for trade-weighted foreign government spending, government consumption vs. investment

Note: x-axis: years, y-axis: percentage points; average (non-cumulative) response in a euro area country to a trade-weighted increase in government spending in the other euro area countries, 5-variables PVAR, 2 lags, LSDV, 68% confidence bands from Monte-Carlo simulations, government consumption (left) and government investment (right).

2.5.4 Robustness checks

We perform a number of robustness checks for the VAR specification, such as the lag structure, the ordering of variables and including additional additional variables. The benchmark results are robust to alternative model specifications and other identification assumptions.

Lag structure. As a first robustness check, we estimate the benchmark model with a different lag structure, i.e. one lag instead of two lags, as selected as the optimal lag structure by the Bayesian information criterion[13]. Average cumulative spillover multipliers are quantitatively close in the two specifications, with an output effect of 0.35 after four years in the one-lag model compared to 0.32 in the benchmark model (see table 2.3). The time profile differs however slightly between the two specifications and the one-lag model shows a slower but more persistent transmission of foreign government spending to domestic GDP. The cumulative multiplier in the alternative specification is 0.25 after two years against 0.42 in the benchmark case. The cumulative domestic multipliers are very close in the two specifications.

Ordering. Next, we change the ordering of the variables in the model. This has an impact on the sequencing of shocks in the structural form of the VAR. In the benchmark model, we assume that trade-weighted foreign government spending is the most "exogenous" variable, i.e. it has an instantaneous impact on all other domestic variables, while domestic variables influences foreign government spending only with a lag. This

[13]We chose the two lag-structure for the benchmark model, as one lag is not enough to control for serial correlation in the residuals

period	benchmark	1 lag	govfor last	incl. reer
0	-0.23	-0.38*	0.00	-0.26
1	0.33	0.07	0.40*	0.30
2	0.42*	0.25	0.36*	0.39*
3	0.39	0.32*	0.31*	0.34
4	0.32	0.35*	0.26*	0.26

Table 2.3: Cumulative multipliers for a €1 increase in trade-weighted foreign government spending

Note: * indicates significance at the 68% significance level, 1972 to 2017

sequencing is the most plausible from an economic point of view (domestic shocks do not affect directly fiscal action in another country). We also test for an alternative ordering where foreign government spending (our main variable of interest) is ordered last and can hence react instantaneously to all domestic variables. One could imagine this sequencing a situation where fiscal action in a country reacts instantaneously to a deterioration of the macroeconomic outlook in another country (for example, a large neighboring country). The results in column three in table 2.3 show that changing th eordering of variables has only a minor impact on the results, which are close to the benchmark specification. The maximum spillover effect is 0.40 after one year in the alternative model, compared to 0.42 after two years in the benchmark model.

Adding exchange rates. Finally, we augment the benchmark model with the real effective exchange rate. In a standard open-economy framework, a fiscal expansion in one country would put upward pressure on prices of this country, leading to an appreciation of its real effective exchange rate (reer) and an increase of its imports as foreign goods become relatively cheaper. In our setup the spending shock comes from multiple source countries, which is why we would have to capture the average reaction of the reer of the different source countries vis-à-vis the destination country. This is identical to using directly the destination country's reer relative to the rest of the 10 euro area countries. Therefore, we rely on the reer of the destination country relative to the former EU-15[14], for which data is readily available in Ameco. We order this variable last in the model since we expect the reer to adjust slowly to government spending shocks. Figure 2.8 shows the reaction of the destination country's reer to a trade-weighted foreign government spending shock. The reer only slightly depreciates, as we would expect from economic theory[15]. The reaction is however not significant at the 68% level. This small and insignificant effect, over the short period that we examine, corresponds to what could be expected for countries in a currency union where the nominal exchange rate is fixed (and relative prices

[14]In addition to the euro area countries in our panel, this includes also Luxembourg, Denmark, the United Kingdom and Sweden

[15]The reer of the source countries would appreciate on average against the destination country.

adjust only slowly). Consequently, the output spilover effects are almsot unaffected by the inclusion of the reer and close to the benchmark model (0.39 after two years and 0.26 after four years against respectively 0.42 and 0.32 in the benchmark model).

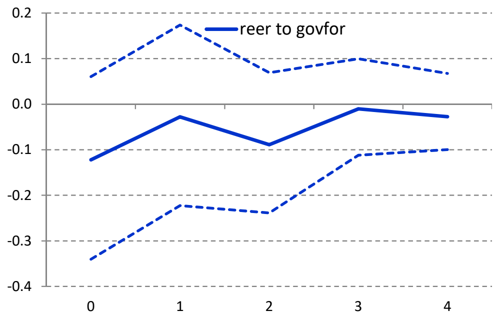


Figure 2.8: Non-cumulative IRF for reer for trade-weighted foreign government spending

Note: Note: x-axis: years, y-axis: percentage points; average (non-cumulative) response in a euro area country to a trade-weighted increase in real government spending in the other euro area countries, 6-variables PVAR (incl. reer), 2 lags, LSDV, 68% confidence bands from Monte-Carlo simulations, 1972 to 2017.

2.6 Conclusions

In this chapter, we provide new empirical estimates for fiscal spillovers in the euro area for the period from 1972 to 2017 using a Panel VAR approach. Spillovers are expressed as destination spillovers, which measure the spillovers from a government spending shock in all but one euro area country on the recipient country. Our empirical results confirm the existence of positive fiscal spillovers within euro area countries. A €1 increase in government spending in the other euro area countries is estimated to affect domestic GDP on average by a cumulative €0.33 after one year, €0.42 after two years (the latter effect being significant at the 68% significance level), before declining to €0.32 after four years (not significant at the 68% significance level). This compares to a cumulative domestic spending multiplier of €0.71 after one year, €0.87 after two years, before decreasing slightly to €0.84 after four years according to our estimates. This results lie within the magnitude of previously estimated spillovers, which, however, are not necessarily directly comparable due to different identification and caluclation schemes and different time and country panels. Our results are robust to alternative model specifications and identification assumptions.

Our results show that the magnitude of the spillovers depends crucially on the size and the degree of openess of the countries where the fiscal shock originates. Fiscal spillovers are larger when originating in more open and closely integrated economies and smaller for more closed economies. The difference is relatively large. For example, fiscal action in one of the four larger economies of the euro area (Germany, France, Italy and Spain) only leads to a cumulative average spillover effect of 0.15 (not significant at the 68% significance level) after four years, whereas the smaller, more open euro area economies can generate a cumulative average spillover effect of up to 0.6 after four years. Other factors also play a role in the size of the spillovers. We show that the size and persistency of the spillover effects have increased slightly over the period under consideration. This may be due to an increase in trade integration over the period, a smaller monetary policy reaction due to the introduction of the single currency or other business cycle factors.

Fiscal spillovers effects, though positive, fall short in size of the stabilising properties of domestic fiscal policy in the euro area. This has important policy implications. A well-designed fiscal capacity might be better suited than coordinated fiscal policies in delivering the right amount and the right timing of fiscal action in the euro area.

Chapter 3

Explaining and Forecasting Euro Area Inflation: the Role of Domestic and Global Factors[1]

3.1 Introduction

Understanding inflation developments and predicting them accurately is of paramount importance for central banks, whose objective is to deliver price stability. From 2013 up to 2019, inflation in the euro area was very low and over-predicted. Forecasts by the Eurosystem and by other institutions were constantly surprised on the downside by inflation developments. According to Ciccarelli and Osbat (2017) and Koester et al. (2021), weak cyclical developments and negative global shocks explain these inflation developments - especially in the first part of this low inflation period. The goal of this chapter is to analyse whether the tools used at central banks are suitable to understand these inflation developments. One of these tools is the Phillips curve, which in its simplest form links inflation to real domestic activity. Augmented with global factors, the Phillips curve has long been an important tool to understand the domestic and global drivers of domestic inflation dynamics and guide the conduct of monetary policy. In this chapter, we examine the explanatory power and the forecast ability of the standard and the augmented Phillips curves for the euro area with respect to different inflation measures (headline and core inflation), different time periods and different predictor variables, especially different global indicators, paying special attention to the last period of low inflation. We distinguish between their explanatory power and their role in forecasting. We address this issue along two dimensions:

First, we re-investigate the goodness-of-fit of the Phillips curve with respect to a

[1] A version of this chapter, co-authored with S. Bereau (Université de Namur and Université de Lorraine) and V. Faubert (Banque de France), has been published as Working Paper Banque de France in February, 2018. The Introduction, Literature Review and Conclusion have been amended for this thesis to reflect developments since 2018.

large set of different domestic and global indicators. We look at more traditional global indicators, such as commodity prices, exchange rates and import prices, and at more broad global indicators, such as global consumer inflation and global economic slack. We find that traditional commodity and import price indicators provide a good identification of the augmented Phillips curve for euro area headline inflation, in contrast to global slack measures as proposed by Borio and Filardo (2007). We cannot identify a significant impact of global factors on core inflation measures in this reduced-form Phillips curve framework. We then study the forecast behaviour of the Phillips curve over different time periods, including the recent episode of low inflation. We focus on the forecast up to one year ahead. We find a lot of time instability in the forecast performance of the Phillips curve for headline inflation against univariate benchmarks. The Phillips curve forecasts perform better during some period than during other periods. The forecast ability of the Phillips curve for core inflation is more stable and provides generally an improvement to univariate benchmarks. In what regards the role of global indicators, we argue that there are important for understanding headline inflation dynamics *ex post*, but their purpose for forecasting is generally limited. Their influence on headline inflation reflects primarily commodity price movements, which are hard to forecast. This is also why global indicators capturing wider global forces, such as global consumer inflation, do not seem to contain additional information for domestic inflation dynamics compared to the more traditional commodity price indicators, such as oil prices. We show however that trade-weighted foreign demand (i.e. the import demand of the euro area trading partners) possesses certain leading properties for euro area inflation and improves the short-term inflation forecast over the analysed period.

Second, we build on the scarce literature on quantile regressions to analyse the Phillips curve relationship and the role of global factors on the entire conditional distribution of inflation. The goal is to understand whether the impact of these variables on inflation is different on specific areas of the conditional distribution, i.e. during low or high inflation periods. Despite the extensive literature on inflation analysis and forecasting, little attention has been paid to the question whether the different indicators carry useful information for other parts of the conditional distribution than the mean and possible non-linearities of the Phillips curve. We find that the inflation process is much more persistent at the left tail of the distribution, i.e. for lower quantiles of inflation. This might explain why mean models have not captured sufficiently the persistency of inflation in the past period of low inflation. In contrast, domestic activity - and to a lesser extent also global factors - are found to have a stronger impact during periods of higher inflation. Turning to forecast considerations, we show that quantile regressions can improve the short-term forecast ability of the Phillips curve during some periods of persistently low inflation (2014-2015).

Since 2021, with the end of the pandemic and the beginning of the Ukraine war, inflation dynamics have rebounded and inflation has strongly increased in the euro area and elsewhere. The question of the influence of global and domestic factors on inflation developments and the possibility of predicting them remains however topical (see, for exemple, Nickel et al., 2022 for an analysis of recent inflation drivers in the euro area). Given the complex interactions of domestic and external forces, demand and supply factors, unparalleled policy measures and very heterogenous impacts of the pandemic across sectors, modelling and isolating the different influencing factors in a single equation model like the Phillips curve is even more challening (see Bobeica and Hartwig, 2021 for a discussion of the modelling challenges). However, given the continuing practical relevance of the Phillips curve for understanding and forecasting inflation at central banks, this tool remains important and we think the conclusions of this chapter as well.

The remainder of the paper is organized as follows: Section 3.2 briefly surveys the empirical literature on the Phillips curve and on the quantile regression approach. Section 3.3 analyses the augmented Phillips curve for mean inflation and presents its in-sample and out-of-sample properties. In Section 3.4 we implement some testing and explore the forecast ability of the Phillips curve over time. In Section 3.5, we examine whether the domestic activity variables and global factors selected in the previous sections help in predicting the entire conditional distribution of inflation. We then compare the forecast performance of OLS and quantile regression models to evaluate whether quantile regression techniques can hedge against a bad forecast performance in particular episodes, such as the recent period of subdued inflation. Finally, Section 3.6 concludes.

3.2 Literature review

This chapter is related to three strands of literature: First, we draw on the literature related to the identification and the forecast performance of the Phillips curve. Secondly, we relate to the literature which studies the role of global factors in domestic inflation. Finally, we explore quantile regressions to analyse the entire conditional distribution of inflation.

Phillips curves and its forecast performance. Various forms of Phillips curves have been used to forecast inflation[2]. Stock and Watson (2008) provide an extensive literature review for the U.S. The literature's conclusions on the forecast performance of the Phillips curve strongly depend on the forecast period, the inflation series and the benchmark models. Atkeson and Ohanian (2001) considered a number of standard Phillips curve forecasting models for U.S. inflation and show that none improve upon a random walk benchmark over the period 1984-1999. In contrast, Stock and Watson (2008) argue that Phillips curves can be useful during some periods, such as the late 1990s, and Stock and Watson (2010) show that some indicators perform better than others.

[2]Following Stock and Watson (2008), we call Phillips curves those models which include an activity variable, such as the unemployment rate or the output gap, perhaps in conjunction with other variables, such as external supply shock indicators.

Though comprehensive studies on the forecast performance of the Phillips curve have been undertaken for the U.S., fewer works are available for the euro area. Banbura and Mirza (2013) examine the forecast ability of a wide range of Phillips curve specifications for different measures of euro area inflation (headline, core and GDP deflator) over the period 1994-2011. As Stock and Watson (2008) they find that the results vary substantially with the forecast period, but that on average Phillips curve models improve upon univariate benchmarks, notably for core inflation, even if improvements are typically not large. The unemployment rate/gap and the output growth/gap are often part of their best models and the inclusion of a global indicator also frequently improves the forecast performance. Ciccarelli and Osbat (2017) confirm that the Phillips curve is still relevant for the euro area and conditional forecasts from some Phillips curve specifications capture well the latest episode of low inflation since 2013.

Since the publication of this chapter in 2018, several contributions analysing the performance of the Phillips curve in the euro area have been added to the literature, focusing mostly on the low inflation episode up to 2019 (see, for exemple, Banbura and Bobeica, 2023; Koester et al., 2021; Moretti et al., 2019 or Ball and Mazumder, 2019). In the closest contribution to ours, Banbura and Bobeica (2023) find, based on data up to 2018, that some Phillips curve specifications for the euro area, including notably a domestic output gap measure, outperform the forecast of a univariate benchmark, while external variables do not bring forecast gains. They also conclude that model averaging over a wide range of Phillips curve models[3] is helpful for forecasting. Overall, the different papers agree that the forecast performance of the augemented Phillips curve is not stable over time and some individual indicators work better at certain times than at others.

The role of global factors for domestic inflation. While the role of external price pressure indicators, such as commodity prices, for domestic headline inflation is relatively well documented in the literature, a limited number of studies look at the influence of more general global factors, such as global inflation and global economic slack, on domestic inflation rates. This strand of literature argues that domestic inflation is being increasingly sensitive to global economic conditions which might not only affect domestic inflation indirectly, via its effect on import prices and the domestic output gap, but also directly. One explanation is that globalisation has rendered domestic inflation less responsive to domestic capacity constraints, either because a sudden domestic demand shock would rather bolster imports than increase prices, or because exposure to foreign competitors curtails increases in domestic tradable prices (Guerrieri et al., 2010). Other studies emphasize the role of credible monetary policies that stabilized inflation expectations (Mishkin, 2009): with domestic price expectations well anchored, proportionally more of the variation in domestic inflation rates would be explained by global factors. However, the empirical evidence for the influence of global conditions is at best mixed,

[3]This is a a point not analysed in this chapter.

especially with regard to the role of global economic slack.

Borio and Filardo (2007) show the importance of the global output gap as a determinant of domestic inflation in advanced economies, stating that the role of global factors has increased over time. Auer et al. (2017) argue that, as participation in global value chains increases, competition among economies increases, making domestic inflation more sensitive to the global output gap. They conclude that the growth of global value chains is associated with both a reduction of the impact of domestic slack on domestic inflation and an increase in the one of global slack. However, other studies, such as Ihrig et al. (2010) and Mikolajun and Lodge (2016), find no empirical evidence for a significant impact of global slack on domestic inflation in advanced economies. Ciccarelli and Mojon (2010) analyse the influence of a global inflation factor on domestic inflation in OECD countries and conclude that models including a measure of global inflation consistently outperform univariate benchmarks. This is confirmed by Medel et al. (2014), who conclude that global inflation improves the inflation forecast for headline and core inflation. However, the gains in forecast accuracy are modest: among the euro area members in the sample, only Italy and Slovakia achieve reductions in RMSE that are relevant (i.e. higher than 5%). In contrast, Mikolajun and Lodge (2016) argue that, with the exception of commodity prices, there is little reason to augment the standard Phillips curves for advanced economies with global factors once the volatile inflation period of the 1970s-1980s is excluded from the sample. They find that from the mid-1990s onwards, the coefficients of global inflation are insignificant for most OECD countries: global inflation measures are helpful for forecasting domestic inflation during periods of high and volatile inflation (i.e. the 1970-80s), but less so since inflation has receded. Using a richer VAR modelling approach, Bianchi and Civelli (2015) show how global economic fluctuations affect the dynamics of domestic inflation. They conclude that single equation models such as the Phillips curve could easily underestimate the role of global factors in the inflation process. However, they also observe that the foreign output gap does not become relatively more important than the domestic output gap for inflation dynamics.

Some recent papers on the influence of globalisation have provided additional support for the role of global factors in domestic inflation. Notably Forbes (2019) find that adding global variables to Phillips curve models, including the global output gap, and allowing the coefficients to change over time can significantly improve the ability of these models to explain inflation. However, they do not analyse the predictive performance of these indicators. Banbura and Bobeica (2023) confirm our finding that external developments, as reflected in commodity prices, exchange rates, foreign slack or inflation, help to explain domestic inflation developments, however, they are less useful for forecasting considering their high volatility and the diffculties to adequately forecast these factors.

Quantile regressions. While much of the literature focuses on analysing the conditional mean of inflation, it might be also interesting to examine the relationship between inflation and its determinants in other regions of the conditional distribution and to produce forecasts away from the conditional mean. As such, only very few papers explore whether domestic activity variables and global factors are useful for analysing other moments of the conditional distribution of inflation. Much of this scarce literature focuses on the U.S. Tillmann and Wolters (2015) use quantile regressions to examine the persistence of the conditional distribution of U.S. inflation and find evidence for a reduction in persistence on all conditional quantiles over time. More recently, Korobilis (2017) introduces Bayesian model averaging methods into quantile regressions and finds that different macroeconomic and financial predictors are relevant for each quantile of U.S. inflation. As the closest work to ours, Manzan and Zerom (2013) show that economic activity indicators, such as the unemployment rate and housing starts, are useful for forecasting the distribution of U.S. inflation, especially at the left tail of the distribution. To the best of our knowledge, only one paper (Busetti et al., 2015) relies on quantile regressions to analyse the conditional distribution of euro area headline inflation. They find that quantile regressions provides superior forecasts to those from a benchmark univariate trend-cycle model with stochastic volatility over the very short forecast horizon during the 2010-2014 period. The conditional quantile regression approach also allows them to describe the underlying features of the conditional distribution of inflation, with a higher persistency of the inflation process in the lowest quantiles and a higher reactivity of the inflation process to activity variables at higher quantiles.

Since the publication of this chapter in 2018, some additional contributions on "augemented" quantile Phillips curve model have been published that highlight the importance of considering the entire conditional distribution of inflation. Lopez-Salido and Loria (2020) study inflation drivers over the inflation distribution both for the U.S. and the euro area. For the euro area, using data from 1999Q1 to 2017Q4, they show that the upper tail is the more sensitive part of the inflation distribution to domestic slack, while the lower tail responds little. The median and the lower tail of the distribution seem more responsive to relative prices of imported goods than the upper tail of the distribution of inflation. Busetti et al. (2021) show that the response of euro area inflation to domestic demand pressures is overall stronger in the upper quantiles of the distribution, while the reaction to global indicators is more similar across the distribution, using en expectile regression approach with time-varying parameters.

This conclusion, which confirms our analysis that the Phillips curve is potentially flatter in lower inflation regimes and steeper in high inflation regimes, is also relevant for the current high inflation episode. At higher inflation rates, upside risks of inflation to domestic demand pressure gap become more prominent[4]. Harding et al. (2023) further elaborate the idea of a non-linear Phillips curve for explaining the current high inflation developments. They confirm that all shocks transmit stronger to inflation when inflation

[4]It could explain why some recent contributions, for exemple Hobijn et al. (2023) across a sample of 29 advanced countries, find a "steepening" of the Phillips curve in the current high inflation environment.

is surging.

3.3 The augmented Phillips curve for mean inflation

3.3.1 Methods

Econometric specification. We investigate the importance of global factors for euro area inflation by augmenting standard Phillips curves with various global factors. We estimate an aggregate equation for the euro area as a whole, using quarterly data over the period 1996Q3-2016Q4. The model, a backward-looking specification of the Phillips curve[5] including a lagged inflation term, is closely related to the "triangle" model proposed by Gordon (1988) and is similar to the type of models considered in Stock and Watson (2008). Our equation is of the following general form:

$$\pi_t = \alpha + \rho(L)\pi_{t-1} + \beta(L)y_t + \gamma(L)z_t + \varepsilon_t \tag{3.1}$$

where the dependent variable π denotes the quarterly rate of inflation at time t, computed as the first difference in the logarithm of the HICP, y a measure of domestic slack, z a global factor and L are lag polynomials.

Models are estimated by OLS[6], using heteroskedasticity and autocorrelation consistent (HAC) estimates of the covariance matrix to address slight serial correlation in the residuals. The optimal lag order is selected on the basis of the Schwarz (BIC) information criterium. Given the limited time span of our data, the maximum number of lags has been limited to four quarters.

Forecast setting. We conduct pseudo-out-of-sample forecasts similar to the ones produced in Stock and Watson (2008) to evaluate the forecast performance of the augmented Phillips curves. For this, we use information available up to t for the predictor variables to estimate the model and then compute the h-period ahead forecast for π_{t+h} based on the direct method. The estimation is then rolled forward one quarter at a time over a fixed forecast period (40 and 74 quarters in our case) and the forecast exercise is repeated. We compute both the one-quarter-ahead ($h = 1$) and the four-quarter-ahead forecast ($h = 4$). We use final data and disregard revision issues in this paper. Our forecast equation is

[5] Recent studies on euro area inflation suggest that backward-looking Phillips curves fit inflation better than forward-looking Phillips curves (Mikolajun and Lodge, 2016). We also test inflation expectations in hybrid Phillips curve specifications, but generally focus on backward-looking specifications.

[6] Following Mikolajun and Lodge (2016), we also estimate Equation 3.1 by the generalized method of moments (GMM) using lags as instruments to address possible endogeneity problems, in particular in the models with global consumer inflation. However, the J-statistics of the Durbin-Wu-Hausman test do not signal any endogeneity. We hence maintain OLS estimates given the risk of incorrect inferences by using weak instruments in GMM estimates.

hence of the following form:

$$\pi_{t+h} = \alpha + \rho(L)\pi_t + \beta(L)y_t + \gamma(L)z_t + \varepsilon_{t+h} \tag{3.2}$$

where h denotes the forecast period.

Benchmarks. We compare the accuracy of the inflation forecasts of the augmented Phillips curves to those from two benchmark models[7] namely an autoregressive model of order 1 (AR hereafter) and a standard backward-looking Phillips curve (PC hereafter), both computed using the direct method, where the latter has the following form:

$$\pi_{t+h} = \alpha + \rho(L)\pi_t + \beta(L)y_t + \varepsilon_{t+h} \tag{3.3}$$

Forecast evaluation. We use the root mean squared forecast error (RMSE) as metric to compare forecasts from the different models. The RMSE corresponds to the square root of the arithmetic average of the squared differences between the actual inflation rate and the predicted inflation rate. The RMSE for a h-period-ahead forecast corresponds to Equation 3.4, where $\pi_{t+h|t}$ is the pseudo out-of-sample forecast of π_{t+h} made using data up to date t.

$$RMSE(t_1, t_2) = \sqrt{\frac{1}{t_2 - t_1 + 1} \sum_{t=t_1}^{t_2} \left(\pi_{t+h} - \pi_{t+h|t}\right)^2} \tag{3.4}$$

Following Stock and Watson (2008), we also compute (bi)weighted rolling estimates of the RMSE (BRMSE hereafter), which correspond to Equation 3.5. Rolling estimates are based on weighted centred 15-quarters windows: bigger (lower) weights are given to errors close to (far from) the centre of the window. Rolling RMSE help to distinguish during which specific period the forecasts performed best.

$$BRMSE(t) = \sqrt{\frac{\sum_{s=t-7}^{t+7} K\left(\frac{|s-t|}{8}\right)\left(\pi_{s+h} - \pi_{s+h|s}\right)^2}{\sum_{s=t-7}^{t+7} K\left(\frac{|s-t|}{8}\right)}} \tag{3.5}$$

[7]We also checked the forecast performance against a random walk. Given the fact that the random walk does not outperform the AR on our fixed forecast windows (it has a similar forecast ability for headline inflation over the longer forecast period and a lower forcast ability over the short forecast period compared to the AR), we decided to use the AR as benchmark.

where K is the biweight kernel :

$$K(x) = \frac{15}{16}(1 - x^2)^2 I(|x| \leq 1)$$

Sample. Our sample covers data over the 1996Q3 to 2016Q4 period, which corresponds to 82 observations at a quarterly frequency. It includes episodes of important volatility in oil prices (which increased dramatically in 2008 and 2011 before decreasing from 2014 onwards), the Great Recession period, as well as the euro area sovereign debt crisis. These major events might have had altered the link between global factors and domestic inflation. Consequently, we compute RMSE over different forecast periods to make sure our models perform well during different time periods. Our in-sample analysis is based on the entire sample period from 1996Q3 to 2016Q4. Estimation results are provided in Appendix 3.B. Our pseudo-out-of-sample forecast analysis relies on a fixed size rolling window approach. Three different procedures have been adopted:

- Models are estimated on the longest possible time span, using rolling estimation windows of a fixed length of 74 quarters. Hence, the first one-quarter-ahead forecast starts in 2015Q1 and the last one-quarter-ahead forecast ends in 2016Q4. The RMSE are computed for a forecast period of 8 observations ($t_2 - t_1 + 1 = 8$) for $h = 1$.

- Models are estimated using a rolling scheme with a shorter rolling estimation window of 40 quarters. RMSE are computed on a 39-quarters forecast period for $h = 1$ and $h = 4$. Hence, the first one-quarter-ahead forecast starts in 2006Q3, and the first one-year-ahead forecast starts in 2007Q2. The last one-quarter-ahead forecast ends in 2016Q1, and the last one-year-ahead forecast ends in 2016Q4.

- Models are estimated on rolling estimation windows of a fixed length of 40 quarters. RMSE and weighted BRMSE are computed on a 15-quarters forecast period for $h = 1$ and $h = 4$. Hence, the first one-quarter-ahead forecast starts in 2006Q3, and the first one-year-ahead forecast starts in 2007Q2.

Rolling estimates on the relatively short-sized 40-quarters window allow us to investigate the importance of the different predictor variables over a longer forecast period. Estimates on the longer, 74-quarters window assure that the results are not biased by the small size of the 40-quarters window and allow us to examine the forecast performance over the latest period of persistently low inflation.

3.3.2 Data

Dependent variables

We examine the Phillips curve for three measures of consumer price inflation: the euro area headline Harmonized Index of Consumer Prices (HICP), the euro area HICP excluding energy (HEX hereafter) and the euro area HICP excluding food and energy (CORE hereafter). We convert monthly HICP data to quarterly data by computing the average value for the three months in the quarter[8]. We seasonally adjust the quarterly indices using the X-12-ARIMA procedure. We focus on the results on headline inflation in the main text in line with the ECB target of overall price stability.

Regressors

Domestic slack. We test different measures of domestic slack for the euro area, namely: (*i*) the unemployment rate; (*ii*) the output gap; (*iii*) the unemployment gap; and (*iv*) the Industrial Production Index (IPI). Most of the measures are stationary and are introduced in levels. The IPI is tested both in level and in variation. To ensure robustness, we rely on different measures of the euro area output gap, derived from statistical filters and from the production function approach. We test: (*i*) the output-gap computed as the log-difference between actual and potential GDP, the latter being measured by means of a Hodrick-Prescott filter; (*ii*) the output gap computed by the European Commission; as well as (*iii*) the output gap computed by the ECB for the staff macroeconomic projection exercise. As far as output gap measures are annual, we used cubic splines techniques to interpolate annual figures into quarterly ones. Our in-sample and out-of-sample analyses show the best performance for models using the output gap computed by the ECB, which we use hence as our benchmark measure. We also report results based on the unemployment rate as the unemployment rate has the advantage of being less affected by revisions than the output gap.

Global factors. "Triangle" models of the Phillips curve traditionally capture external cost-push factors via import prices or commodity prices. We test a large number of these traditional used external factors including: (*i*) changes in oil prices; (*ii*) changes in the price of other commodities; (*iii*) changes in the euro area bilateral and effective exchange rates; and (*iv*) changes in euro area import prices, which can influence domestic inflation via the price of imported commodities, the price of imported final consumer goods as well as the price of imported intermediate goods. Concerning the latter, we consider three different indicators of import prices: (*i*) the extra-euro area import deflator for goods and services; (*ii*) the relative import deflator, i.e. the ratio of the extra-euro area import deflator to the GDP deflator; and (*iii*) competitors' prices on the import side[9].

In order to capture the growing international integration of goods and labour markets and the wider propagation of global cost shocks, we furthermore test indicators proposed

[8] Though year-on-year inflation has no seasonal pattern, using year-on-year rates may introduce a moving average component to inflation. Annual inflation measured by year-on-year rates is approximately equal to the sum of quarterly log HICP differences. As a result, using year-on-year rates can complicate econometric inference, with autocorrelated residuals. We therefore rely on seasonally adjusted quarter-on-quarter rates in our estimations.

[9] The euro area competitors' prices are computed by the ECB as a weighted average of trading partners' export prices (Hubrich and Karlsson, 2010).

in the recent literature such as global consumer inflation and global economic slack. As a measure of global consumer inflation, we succesively consider: (*i*) a simple average of cross country inflation rates[10]; and (*ii*) a weighted average of cross country inflation rates[11], both for the total CPI and the CPI excluding energy and food (CORE). For the global economic slack, we use different measures of the output gap and the unemployment rate. For the output gap, we consider: (*i*) output gaps computed as the difference between actual and potential GDP, the latter being computed by means of a Hodrick-Prescott filter; and (*ii*) output gaps computed by the IMF. Different weighting schemes are applied to compute the output gap of various groups of countries: (*i*) cross-country simple averages; and (*ii*) weighted averages, taking relative GDP as weights. We consider several groups of countries to compute our global measures: the US, the OECD excluding members from the euro area, major advanced economies (i.e. the U.S., the U.K., Japan and Canada), major emerging and advanced economies excluding members from the euro area (world hereafter) and major emerging market economies.

We also test the euro area trade-weighted foreign demand index (FDR)[12]. This trade-weighted indicator of global demand is likely to reflect global demand-related price pressures that have an impact on the euro area better than non-trade weighted indices. Details regarding the variables and their transformations are provided in Appendix 3.A.

Inflation expectations. We use two measures of inflation expectations in the hybrid specifications of the Phillips curve: (*i*) a survey-based measure for households from the monthly European Commission survey; and (*ii*) a forecast-based measure from the Consensus forecast (more precisely, the one-quarter-ahead and the four-quarter-ahead Consensus forecast).

3.3.3 Results

In-sample evaluation

In this section, we analyse the in-sample fit of the Phillips curve augmented with the different global factors. Our results show an important role for commodity prices, import prices and global consumer inflation for headline inflation, when entered in a contemporaneous relationship with inflation. The coefficients of these global factors are statistically significant and positive for estimations over the entire sample from 1996Q3 to 2016Q4 (see Appendix 3.B). They strongly improve the in-sample fit compared to the two benchmark

[10] Mikolajun and Lodge (2016) note that a simple average closely follows a common factor of global inflation rates. We hence use the simple average as a proxy for a common global factor in our estimations.

[11] Country weights are computed by the OECD and are based on the previous year's private final consumption expenditure of households and non-profit institutions, expressed in purchasing power parities (PPP).

[12] The euro area foreign demand index computed by the ECB (Hubrich and Karlsson, 2010) corresponds to the geometric average of the real imports of the trading partners of the euro area: real imports of goods and services are weighted by the share of a given trading partner in the euro area total exports.

models with an adjusted R2 of close to 0.60. The results are robust to different measures of domestic slack in the augmented Phillips curve, such as the output gap, shown in Table 3.1, and the unemployment rate.

Dependent variable	HICP		HEX		CORE	
Model	Sign.	Adj. R2	Sign.	Adj. R2	Sign.	Adj. R2
AR		0.20		0.38		0.33
PC-OG		0.23		0.46		0.44
OG+Oil price		0.60	GF insign.	0.45	GF insign.	0.43
OG+Non-energy commodities		0.37	GF insign.	0.46	GF insign.	0.43
OG+Import prices		0.56	GF insign.	0.48	GF insign.	0.43
OG+OECD CPI (weight.)		0.59		0.47	GF insign.	0.43
OG+US CPI		0.60		0.47	GF insign.	0.43
OG+FDR		0.41	GF insign.	0.46	GF insign.	0.44
OG+FDR (lag 4)	GF, OG insign.	0.20		0.45		0.44
OG+OG adv. econ. (lag 1)	GF, OG insign.	0.23	GF insign.	0.46	GF insign.	0.43
OG+OG US (lag 1)	GF, OG insign.	0.23	GF insign.	0.45	GF insign.	0.44
OG+OECD core CPI (weight.)	GF, OG insign.	0.23	GF insign.	0.45	GF insign.	0.43
OG+US core CPI	GF, OG insign.	0.23	GF insign.	0.45	GF insign.	0.43
OG+Consumer inflation exp.	OG insign.	0.30		0.53		0.47
OG+Consensus	OG insign.	0.55		0.55		0.47

Adjusted R2 for Phillips curves with the output gap estimated over the full sample period 1996Q3-2016Q4. "GF insign." stands for an insignificant coefficient of the global factor and "OG insig." for an insignificant coefficient of the domestic slack measure at a 10% significance level. Estimation details are reported in Appendix 3.B.

Table 3.1: Adj. R2 for the augmented Phillips curves and benchmark models

The augmented Phillips curve with oil prices and the one with global consumer inflation[13] perform equally well in-sample, closely followed by the model with import prices[14]. The coefficients of global core inflation measures[15] are not statistically significant in the augmented Phillips curve for headline inflation. This illustrates that the significance of global consumer inflation should principally reflect the role of commodity prices. The coefficients of the different global slack measures are not statistically significant from zero for estimations performed over the full sample, see for instance the results for the GDP-weighted output gap of advanced economies (excluding the EA) and the US output gap provided in Table 3.1. These results contrast with Borio and Filardo (2007) and Auer et al. (2017), which show a positive and increasing role of global slack measures in domestic inflation rates. They are closer in line with Mikolajun and Lodge (2016), which find that measures of global economic slack are rarely significant (and positively related to domestic inflation) in Phillips curves estimates for the G7 economies. We only find a positive and significant relationship between the global output gap and domestic inflation rates during the small time period of the Great recession (2008-2010), according to rolling window

[13]The weighted OECD headline CPI excluding the EA as well as the US headline CPI.

[14]We tested different import price indicators but only show results for the extra-euro area import deflator here, which has the best in-sample fit.

[15]The weighted OECD core CPI excluding the EA as well as the US core CPI.

estimates (see Figure 3.1). But even during this short period, the models with import prices or global consumer headline inflation fit domestic inflation data much better. The coefficient of the global output gap also loses its significance when it is added as a second global factor next to the more traditional global factors. Hence, we find little evidence for augmenting the Phillips curve with (non-trade-weighted) global economic slack measures, once domestic slack and more direct measures of global price pressures are taken into account. This conclusion is robust to using alternative measures of global economic slack (i.e. the unemployment rate or output gap estimates derived from statistical filters) and different lags structures.

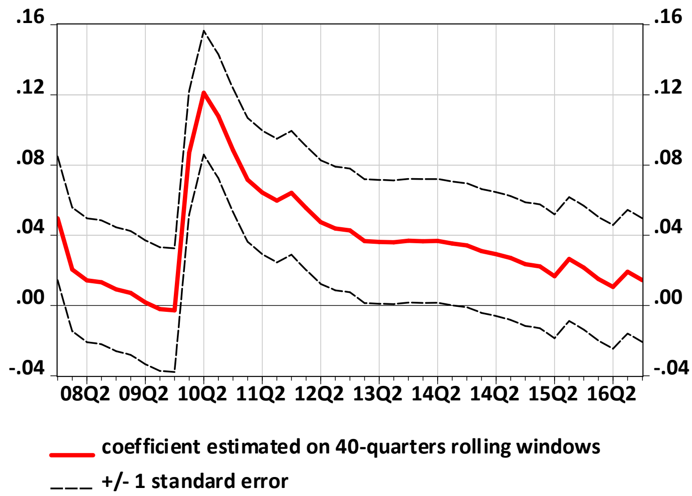


Figure 3.1: Rolling coefficient γ of global output gap in an augmented backward-looking Phillips curve

Note: The initial estimation sample covers 40 quarters from 1996Q3 to 2006Q2. Coefficients are rolled forward one quarter at a time. The model estimated corresponds to Model M8: $\pi_t = \alpha + \rho_l \pi_{t-l} + \beta og_{t-1} + \gamma ogadv_{t-1} + \varepsilon_t$ with π dlog of the headline HICP, og the lagged domestic output gap, $ogadv$ the lagged output gap of advanced economies excluding the euro area.

We find that the coefficient of trade-weighted foreign demand is statistically significant and positive in the augmented Phillips curve, even if the overall fit of the model is lower than the one of our best performing models. It seems that trade-weighted import demand has a more important influence on euro area prices than the general global demand situation, as reflected by non-trade weighted global slack measures. Foreign demand has an additional advantage compared to the commodity price or import price indicators that it seems to possess certain leading properties for domestic inflation. Its coefficient remains statistically significant even when entered with lags, in contrast to the coefficient estimates of other global factors. The hybrid Phillips curve including survey-based or forecast-based measures for inflation expectations explains inflation generally less well than the backward-looking Phillips curve augmented with global factors.

Core inflation. The importance of global factors in the augmented Phillips curve for the two core inflation measures (HICP excluding energy and HICP excluding energy and food) is considerably reduced. The coefficients of the different global factors are generally no longer statistically significant for estimations performed over the full sample from 1996Q3 to 2016Q4, except for a weak significance of the OECD CPI measure (excluding the EA) and the US CPI measure in the Phillips curve for the HICP index without energy (see Table 3.1). The global factors also do not improve the fit compared to the benchmark Phillips curve. This also applies to global core inflation measures[16], which do not show a significant impact on domestic core inflation. Yet, we would not conclude from these results that global factors do not influence domestic core inflation at all. The impact should however be more gradual and dispersed than for the non-core elements, which are strongly driven by commodity price movements, which makes it difficult to identify it in a significant manner in a reduced-form Phillips curve type of model. For example, Chatelais and Schmidt (2017) find a significant impact of import prices on the core element of manufactured goods when relying on models which allow for a more gradual impact over time (error correction models, VAR). Rolling window estimates over a 40-quarters period suggest nevertheless that the impact of global factors on core inflation has increased lately. The coefficient turned significant over the latest period in rolling window estimates (see Figure 3.2), which might reflect the increasing share of imports in core consumer goods. As regards inflation expectation, we find that the inflation expectation measures (for total inflation) enhance the fit of the Phillips curve for core inflation.

Forecast performance

To understand the role of global factors for forecasting inflation, we run the out-of-sample forecast setup outlined in Section 3.3.1 and compare the root mean squared errors (RMSE) of the augmented Phillips curves to our two benchmarks. Table 3.2 shows relative RMSE for the standard and the best performing augmented Phillips curve against the AR. Table 3.3 provides relative RMSE of the augmented Phillips curves against the standard Phillips curve. We report results for two forecast horizons, the one-quarter-ahead and the one-year-ahead forecast. The forecast comparison reveals that the standard Phillips curve for headline inflation only outperforms the AR for the one-quarter-ahead forecast during the most recent period but not during the longer forecast period. For this short period, the improvements in forecast accuracy are noticeable with RMSE reductions of 7% for the standard Phillips curve with the output gap, 6% for the one with the unemployment rate and 14% for our preferred augmented Phillips curve with the output gap and the foreign demand index.[17] For the longer forecast period of the last ten years, we only find a small

[16]The weighted OECD core CPI excluding the EA as well the US core CPI.

[17]Phillips curve with other global factors, such as oil prices or import prices, also achieve reductions in RMSE compared to the AR over the short forecast sample. Results are not shown here as this is basically due to the inclusion of the output gap. These other global factors do not improve the forecast compared to the standard Phillips curve or in a model without the output gap in comparison to the AR.

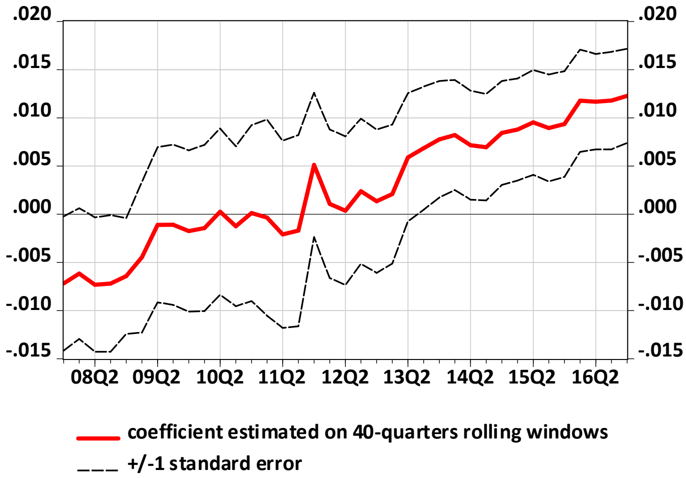


Figure 3.2: Rolling coefficient γ of import prices in an augmented backward-looking Phillips curve for core inflation

Note: The initial estimation sample covers 40 quarters from 1996Q3 to 2006Q2. Coefficients are rolled forward one quarter at a time. The model estimated corresponds to Model M3: $\pi_t = \alpha + \rho_l \pi_{t-l} + \beta og_{t-1} + \gamma z_t + \varepsilon_t$ with π dlog of the HICP excluding energy and food, og the lagged domestic output gap, z the extra-EA import deflator.

improvement in forecast accuracy of 4% for the specification with the output gap and the foreign demand index. No improvements in the forecast performance against the AR are achieved for the one-year-ahead forecast. This result that the forecast performance of the Phillips curve against simple univariate benchmark is episodic and depends on the time period confirms the results for the U.S. in Stock and Watson (2008).

Table 3.3 shows that global factors generally do not improve the forecast ability for headline inflation compared to the standard benchmark Phillips curve. While commodity prices, import prices or global consumer inflation help to explain domestic inflation rates *ex post*, they are less powerful *ex ante* for forecasting purposes. This is even true for the very short next-quarter forecast horizon. Likewise, global core inflation measures or global slack measures do not lead to a better forecast compared to the benchmark Phillips curve. There are only two global indicators, the non-energy commodity price index and the trade-weighted foreign demand index, that achieve reductions in RMSE that are higher than 3%, both over the shorter (2015Q1-2016Q4) and the longer forecast sample (2006Q3-2016Q1) for the one-quarter-ahead forecast. The foreign demand index leads to the biggest gain in forecast accuracy of about 8% for the Phillips curve with the output gap. The usefulness of the foreign demand index is confirmed in the specification which uses the unemployment rate instead of the output gap as a measure of domestic slack. This leads us to conclude that trade-weighted import demand (the foreign demand index) is a useful leading indicator for the short-term domestic inflation forecast. On the contrary, we do not find global consumer inflation to be useful for the domestic inflation

	Dependent variable	Headline HICP			HEX			CORE		
	Estimation window (obs.)	74	40		74	40		74	40	
	Forecast period (obs.)	8	39		8	39		8	39	
	Forecast horizon	h=1	h=1	h=4	h=1	h=1	h=4	h=1	h=1	h=4
	Model									
AR	Autoregressive model	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
PC-OG	PC(OG)	0.93	1.05	1.19	0.82	1.03	1.35	0.77	0.93	1.26
PC-U	PC(UR)	0.94	1.05	1.17	0.94	1.02	1.27	0.93	0.92	1.10
M7	OG+FDR	0.86	0.96	1.18	0.82	1.08	1.22	0.78	0.98	1.11
M12	OG+Consensus	0.93	1.09	1.25	1.12	1.00	1.36	0.90	0.90	1.33

Ratios below 1 signify a lower RMSE for the Phillips curve compared to the AR. Models are estimated on rolling windows of a fixed size of 74 quarters and RMSE are computed over 8 observations for $h = 1$ (first column for each inflation series). Models are estimated on rolling windows of a fixed size of 40 quarters and RMSE are computed over 39 observations for $h = 1$ and $h = 4$ (second and third column for each inflation series). Grey shaded cells highlight situations in which the achieved reductions in the RMSE is at least 3%. OG stands for output gap and UR for unemployment rate.

Table 3.2: RMSE ratios between Phillips curves and the autoregressive model

forecast as in Ciccarelli and Mojon (2010), nor global slack measures as in Borio and Filardo (2007) and Auer et al. (2017). None of the global indicator improves the forecast ability for the one-year-ahead forecast. The hybrid Phillips curve specification including inflation expectations also does not outperform the standard benchmark Phillips curve according to this forecast exercise.

Core inflation. Regarding the out-of-sample forecast performance of the Phillips curves for core inflation (HICP without energy and HICP without energy and food), we find that they outperform the AR for the one-quarter-ahead forecast, both for the shorter forecast sample and, in the case of the HICP index without energy and food, also for the longer forecast sample (see Table 3.2). Average gains in forecast accuracy are higher than for headline inflation, with reductions in the RMSE of up to 8% for the HICP index excluding energy and 23% for the one excluding energy and food. This conclusion is robust to different indicators of domestic slack, such as the output gap and the unemployment rate. The specification with the output gap performs however slightly better than the one with the unemployment rate in this pseudo out-of-sample setting.[18] As could be expected from the limited significance of global factors in the in-sample analysis, global factors generally do not improve the forecast ability of the standard benchmark Phillips curve (see Table 3.3). An exception is the specification with the non-energy commodity price index (Model 2), which achieves small reductions in the RMSE of 3% for the one-quarter-ahead forecast for the most recent forecast period. We find more important reductions in

[18]The unemployment rate benefits however from the fact that it is only little revised, while the output gap estimates can be revised quite substantially, which might give an advantage to the unemployment rate in a real-time forecast environment.

	Dependent variable	Headline HICP			HEX			CORE		
	Estimation sample (obs.)	74	40		74	40		74	40	
	Forecast period (obs.)	8	39		8	39		8	39	
	Forecast horizon	h=1	h=1	h=4	h=1	h=1	h=4	h=1	h=1	h=4
	Model	Backward-looking Phillips curve with the OG								
PC-OG	PC(OG)	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
M1	Oil price	1.01	1.00	1.05	1.02	1.17	1.06	1.08	1.08	1.01
M2	Non-energy commodities	0.92	0.96	1.01	0.97	1.03	1.03	0.97	1.03	1.02
M3	Import prices	1.00	1.00	1.04	1.00	1.08	1.04	1.04	1.06	0.99
M4	OECD CPI (simple)	1.04	0.98	0.99	1.00	1.04	1.01	1.02	1.04	1.01
M5	OECD CPI (weight.)	1.01	1.07	1.08	1.05	1.09	1.12	1.01	1.13	1.06
M6	US CPI	1.00	1.05	1.11	1.04	1.09	1.09	1.01	1.18	1.01
M7	FDR	0.92	0.92	0.99	1.01	1.05	0.90	1.01	1.05	0.88
M8	OG adv. econ.	1.00	1.02	1.01	1.01	1.11	1.02	1.00	1.14	1.05
M9	OECD core CPI (w.)	1.00	1.01	1.02	1.02	1.01	0.97*	1.02	1.00	0.97*
M10	US core CPI	1.00	1.02	1.05	1.05	1.02	0.96	1.03	1.02	0.97*
		Hybrid Phillips-curve with the OG								
M11	Cons. inflation exp.	1.00	1.02	1.00	0.99	0.97	1.04	1.01	1.02	1.11
M12	Consensus	1.00	1.04	1.05	1.37	0.97	1.01	1.16	0.97	1.06
		Phillips-curve with unemployment								
PC-U	PC(UR)	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00	1.00
M13	FDR	0.95	0.93	0.98	0.98	1.11	0.89	0.99	1.08	0.87
M14	Consensus	0.99	1.04	1.11	1.17	0.97	1.05	0.97	0.96	1.09

Ratios below 1 signify a lower RMSE for the augmented Phillips curve compared to the benchmark Phillips curve. Models are estimated on rolling windows of a fixed size of 74 quarters and RMSE are computed over 8 observations for $h = 1$ (first column for each inflation series). Models are estimated on rolling windows of a fixed size of 40 quarters and RMSE are computed over 39 observations for $h = 1$ and $h = 4$ (second and third column for each inflation series). Grey shaded cells highlight situations in which the achieved reductions in the RMSE is at least 3%. OG stands for output gap and UR for unemployment rate. An asterisk marks the case where the coefficient of the global factor has an unexpected (i.e. negative) sign compared to what could be expected from economic theory.

Table 3.3: RMSE ratios between augmented Phillips curves and the benchmark Phillips curve

the forecast error of around 10% for the models augmented with the foreign demand index (Model 7 and 13) for the four-quarter-ahead forecast. Similarly, we also detect smaller improvements in the forecast ability of the Phillips curve augmented with the two global core measures (Model 9 and 10) for the four-quarter-ahead forecast. These models are however unsatisfactory from an economic point of view, as the coefficient estimates (of the fourth lag) of the global core indices have a negative sign (which is non-significant over the full 1996Q3-2016Q4 sample), while economic theory would rather suggest a positive relationship between global and domestic inflation (see also Mikolajun and Lodge (2016)). This renders the interpretation of these forecast results challenging. In contrast to the results for headline inflation, we also find that inflation expectation measures in hybrid Phillips curve specifications contain useful information for forecasting core inflation.

3.4 Robustness analysis

3.4.1 Forecast performance over time

Stock and Watson (2008) highlight the unstable forecast behaviour of the Phillips curve for the U.S. over time. We want to verify this result for the euro area and compute forecast errors of the Phillips curve over rolling forecast samples of 15 quarters. Figure 3.3 shows these rolling forecast errors of the standard Phillips curve and some of our preferred augmented Phillips curve specifications for the one-quarter-ahead forecast for the period from 2006Q3 to 2016Q4[19]. It reports relative BRMSE (i.e. weighted RMSE) of the different models compared to the AR benchmark. The BRMSE allows to distinguish the period during which the forecast performed better by assigning bigger weights for forecast errors in the centre of the forecast window. The figure shows that the standard Phillips curve improves in its forecast ability compared to the AR from 2012 onwards (i.e. over the forecast sample 2009-2012). This is the period where the euro area output gap deteriorated sharply. Its forecast ability increases further during the euro area debt crisis and surpasses the forecast performance of the AR at the end of our sample, i.e. for forecasts from mid-2012 to 2016. The augmented Phillips curve with trade-weighted foreign demand (Model 7) works particularly well at the beginning of the forecast sample (i.e. during the Great recession period 2006-2010 where global demand fell significantly) and also at the end of the sample. The model with oil prices (Model 1) improves in its forecast ability compared to the AR - and also to the standard Phillips curve - during the period of the strong fall of oil prices (i.e. over the forecast sample 2012-2015). This fluctuating pattern of the forecast error confirms that the forecast performance of the Phillips curve for the euro area is unstable and depends strongly on the forecast sample. During some time periods domestic activity variables and global factors provide however a clear forecast advantage to univariate benchmarks such as the AR. This seems to be

[19] The date on the time axis of Figure 3.3 represents the end of the rolling 15-quarters forecast sample.

especially the case when we observe strong movements in the predictor variables.

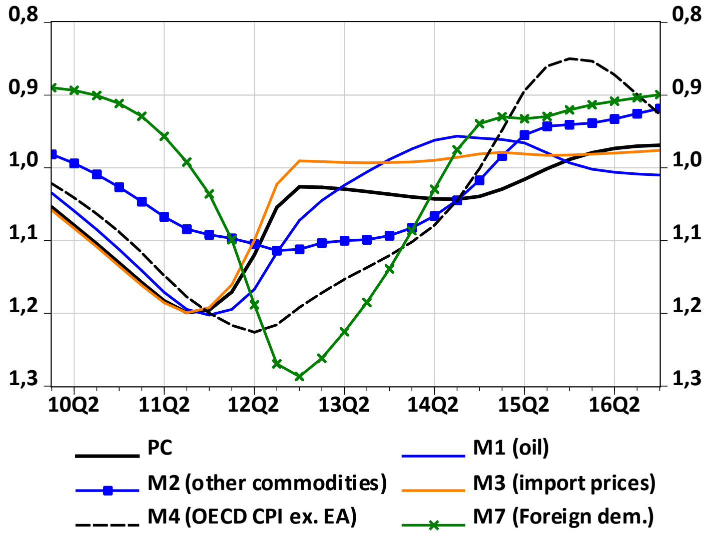


Figure 3.3: Relative BRMSE between Phillips curves and the AR

Note: Relative BRMSE are computed as the ratio of the BRMSE of the standard Phillips curve (PC) and the augmented Phillips curves (Model 1, 2, 3, 4 and 7) to the AR, realized over 15-quarters rollowing forecast windows for the one-quarter-ahead forecast $h = 1$. The date on the time axis represents the end of the 15-quarters forecast window. A smaller (higher) relative BRMSE signifies a better (worse) forecast performance. Scales are inverted.

Giacomini and Rossi (2010) provide a formal test, the so called fluctuation test, to examine whether two models have an equal forecast performance at each point of time of the rolling subsamples. We implement the fluctuation test for the standard Phillips curve to establish more formally whether it outperformed the AR during one of the 15-quarters rolling forecast periods over the 2006q3 to 2016Q4 sample for the one-quarter-ahead forecast. The test is based on the standardised local relative RMSE and we implement it as one-sided test with Clark-West critical values (see Section 3.4.2 for a discussion). The test results, which are shown in Appendix 3.B, broadly confirm the results from the graphical analysis: (i) the test statistic shows a strongly fluctuating pattern and signals an unstable relative forecast performance of the two models; and (ii) the relative performance of the standard Phillips curve increases strongly towards the end of the sample (i.e. for the 2013-2016 period); however, it does not significantly outperform the AR.[20]

[20]Smaller rolling forecast windows of 8-quarters lead to a rejection of the null hypothesis of equal predictability of the two models at the end of the sample, meaning that the Phillips curve outperforms the AR over the last two years of the forecast sample, which is in line with the results in Section 3.3.3.

3.4.2 Test

We perform pairwise Clark-West tests [Clark and McCracken (2001) and Clark and West (2007)] (CW hereafter) to evaluate whether the small gains in forecast accuracy of the Phillips curve reported in Section 3.3.3 are statistically significant. CW propose this test for forecast comparisons in the case of nested models in order to control for the noise in the forecast evaluation coming from additional parameter estimates. The CW-test hence evaluates whether additional parameters - such as the domestic activity variable and the global factors in the Phillips curve - improve the accuracy of the forecast compared to the more parsimonious benchmark. We use the AR as benchmark for the standard Phillips curve and the standard Phillips curve as a benchmark for the augmented and hybrid Phillips curve specifications. We concentrate our analysis on the longer, 39-quarters forecast sample.

Defining the loss-forecast differential given data up to time t for horizon h as:

$$c_{t(h)} = \left(\pi_{t+h} - \pi_{1,t+h|t}\right)^2 - \left[\left(\pi_{t+h} - \pi_{2,t+h|t}\right)^2 - \left(\pi_{1,t+h|t} - \pi_{2,t+h|t}\right)^2\right]$$

with indices 1 and 2 standing, respectively, for the parcimonious benchmark and the larger model. The null hypothesis of the CW-test is then given by:

$$H_0 : \mathbb{E}[c_{t(h)}] = 0$$

vs. the alternative:

$$H_1 : \mathbb{E}[c_{t(h)}] > 0$$

The test is one-sided, as we are trying to establish whether model 2 is superior to model 1, or whether we cannot distinguish upon their forecast performance. In the latter case, model 1 is said to encompass model 2. Under the null hypothesis, the test statistic is defined as follows:

$$\mathsf{CW}_h = \frac{\bar{c}_h}{\sqrt{\sigma^{2,\mathsf{LR}}_{\bar{c}_h}/T}} \tag{3.6}$$

where $\bar{c}_h$ stands for the sample mean of $c_{t(h)}$ and $\sigma^{2,\mathsf{LR}}_{\bar{c}_h}$ is a consistent estimate of the asymptotic long-run variance of $\sqrt{T}\cdot\bar{c}_h$. Given the small sample size in the rolling window estimates, we compare the test statistic to critical values from a Student-t distribution instead of asymptotically normal critical values, as proposed in Clark and West (2007). The test requires that multistep forecasts are conducted with the direct method, as it is the case in our forecast exercise.

Table 3.4 reports the $\mathbb{P}$-values of the CW-test. The standard PC significantly outperforms the AR for core inflation (HEX and CORE), but not for headline inflation. The

	Dependent variable	Headline HICP			HEX			CORE		
	Benchmark	AR	PC		AR	PC		AR	PC	
	Forecast horizon	h=1	h=1	h=4	h=1	h=1	h=4	h=1	h=1	h=4
	Model									
PC-OG	PC(OG)	0.51			0.07			0.00		
M7	OG+FDR	0.01	0.01	0.03	0.12	0.63	0.00	0.02	0.72	0.00
M1	OG+Oil price		0.26	0.96		0.97	1.00		0.68	0.57
M2	OG+Non-ener. comm.		0.07	0.82		0.58	0.99		0.59	0.58
M5	OG+OECD CPI		0.93	0.37		0.92	0.99		0.82	1.00
M6	OG+US CPI		0.88	0.80		0.94	0.99		0.74	0.61
M9	OG+OECD core CPI		0.24	0.00*		0.56	0.01*		0.43	0.02*
M10	OG+US core CPI		0.61	0.08*		0.90	0.02*		0.99	0.06*
M11	OG+Consumer exp.		0.73	0.40		0.00	0.37		0.02	0.56
M12	OG+Consensus		1.00	0.74		0.03	0.59		0.01	0.99

Models are estimated over rolling windows of a fixed size of 40 quarters and forecast errors are computed over 39 quarters for $h = 1$ and $h = 4$. Grey shaded cells highlight situations in which the null hypothesis that the small nested model performs as well as the larger model is rejected at the 10% level. The standard PC and Model 7 are compared to the AR while the augmented PC and the hybrid PC are compared to the standard PC. An asterisk marks the case where the coefficient of the global factor has an unexpected (i.e. negative) sign compared to what could be expected from economic theory.

Table 3.4: Pairwise Clark-West test ($\mathbb{P}$-values)

Phillips curve augmented with foreign demand (Model 7) is significantly better than the AR both for headline inflation and for core inflation (for the CORE measure). As regards the comparison to the standard Phillips curve, only the augmented Phillips curve with foreign demand provides significantly better forecasts for all three inflation series: over the two forecast horizons for headline inflation and over the one-year-horizon for core inflation (HEX and CORE). The inclusion of non-energy commodity prices also provides a significant, if weak, improvement of the forecast ability of the Phillips curve for headline inflation for the one-quarter-ahead forecast. The CW-test also demonstrates that the global core measures significantly improve the forecast ability of the Phillips curve for core inflation (HEX and CORE). However, as highlighted in Section 3.3.3, the results are driven by negative sign estimates of the global core measures, which complicates the economic interpretation of these results.

3.5 Quantile regressions

In this section, we investigate the relationship between inflation and its determinants on the entire conditional distribution of inflation and not just on the conditional mean. The aim is to understand whether the impact of the dependant variables varies across inflationary regimes and if this can be used for forecasting. For this, we rely on the dynamic quantile regression approach. Our quantile regressions are motivated by Phillips curve arguments from Section 3.3, i.e. they include a domestic slack measure and global factors. The quantile regression approach has the advantage that it allows to model non-linearities or asymmetries in the Phillips curve relationship. There is indeed an intensive discussion in the literature about possible non-linearities in the reaction of inflation to domestic activity (see Musso et al. (2007) and Dolado et al. (2005) for a good overview on non-linearity issues). The quantile regression approach is a relative flexible approach in this sense. It allows for a non-linear reaction of inflation to domestic activity without having a prior assumption of the exact form of the asymmetry.

3.5.1 Methods and specification

Methods. In this paragraph, we briefly review the quantile regression approach, for further details see Koenker (2005).

Let $F_Y(y) := \mathbb{P}[Y \leq y]$ be the cumulative distribution function of a random variable Y. As recalled by D'Haultfoeuille and Givord (2014), for any $0 < \tau < 1$, the τ-th quantile of Y, denoted $q_\tau(Y)$ is defined by: $q_\tau(Y) := \inf\{y : F(y) \geq \tau\}$, which simplifies to the following relationship when the variable of interest Y (as in our case, inflation) is continuous:

$$\mathbb{P}[Y < q_\tau(Y)] = \tau \Leftrightarrow q_\tau(Y) := F_Y^{-1}(\tau)$$

Let $(Y_t)_{t=1,\ldots,T}$ be a set of i.i.d. observations for Y. An intuitive way to provide an estimator for $q_\tau(Y)$, denoted $\widehat{q}_\tau(Y)$, consists in ordering those T observations by ascending order, the τ-th quantile corresponding to the $[T*\tau]$-th observation.[21] Alternatively, it can be shown that the estimator corresponds to the solution of the following minimization problem (Koenker and Bassett, 1978):

$$\widehat{q}_\tau(Y) = \mathsf{argmin}_q \left\{ \sum_{t=1}^{T} \rho_\tau\left(Y_t - q\right) \right\} = \mathsf{argmin}_q \left\{ \sum_{y_t - q \geq 0} \tau |Y_t - q| + \sum_{y_t - q < 0} (1-\tau)|Y_t - q| \right\}$$

with $\rho_\tau(u) = (\tau - \mathbb{I}_{\{u<0\}}) \cdot u$, the loss function with $1_{\{\mathsf{Condition}\}}$ the indicator function that takes value 1 if the condition in brackets is satisfied, 0 otherwise.
For example, for $\tau = 0.5$, the sample median is obtained as the solution to the problem of minimising a sum of equally weighted absolute residuals: $\widehat{q}_{0.5}(Y) = \mathsf{argmin}_q \sum_{t=1}^{T} |Y_t - q|$. For $\tau \neq 0.5$, quantiles other than the median can be obtained by minimising a sum of asymmetrically weighted absolute residuals, i.e. giving differing weights to positive and negative residuals, where the weights are asymmetric functions of τ.

Specifications. Quantile regressions aim to assess how conditional quantiles $q_\tau(Y|X) = F_{Y|X}^{-1}(\tau)$ change with respect to changes in the dependant variables X. As such, and as

[21] $[T*\tau]$ here characterizes the smallest integer being greater or equal to $T*\tau$.

recalled by Koenker (2005), they can be viewed as an extension of classical least square estimation method of conditional mean models, in the way that they provide estimates for a set of conditional quantile functions assuming potentially heterogeneous specifications. We consider the following general quantile specification:

$$q_\tau(\pi_t|X_t) = \alpha^{(\tau)} + \rho^{(\tau)}\pi_{t-1} + \beta^{(\tau)}y_{t-1} + \sum_{l=0}^{Max=4} \gamma_l^{(\tau)} z_{t-l} + \varepsilon_t^{(\tau)} \tag{3.7}$$

with $X_t = (1, \pi_{t-1}, y_{t-1}, z_{t-l})'$ the set of dependant variables as defined in Section 3.3.1. As all our specifications include past inflation, we talk about "dynamic quantile regressions", for which standard techniques and metrics apply. Given the limited size of our sample, we employ resampling techniques (bootstrap) to derive inference, as suggested by Koenker and Xiao (2002). Models are estimated for different quantile orders.[22] Estimation results for the full sample period from 1996Q3 to 2016Q4 are reported in Appendix 3.D. As a measure of goodness-of-fit for quantile regression, we rely on the pseudo R^2 proposed by Koenker and Machado (1999), based on the residual absolute sum of weighted differences (RASW):

$$RASW_\tau = \sum_{y_t - \widehat{q}_{\tau,t}(y) \geq 0} \tau|y_t - \widehat{q}_{\tau,t}| + \sum_{y_t - \widehat{q}_{\tau,t} < 0} (1-\tau)|y_t - \widehat{q}_{\tau,t}|$$

where $\widehat{q}_{\tau,t}$ is the fitted τ-quantile at time t. The pseudo R^2 is constructed as the coefficient of determination in OLS and is defined as:

$$R_\tau^2 = 1 - \frac{RASW_\tau}{TASW_\tau}$$

where $TASW_\tau$ is the total sum of weighted differences and with R_τ^2 ranging between 0 and 1.

3.5.2 Estimation results

In a first step, we compute residual densities from linear OLS for the augmented Phillips curves with the best in-sample fit (respectively Model 1 and 3, each with the output gap and the unemployment rate as a measure of domestic slack) in order to compare it to the standard normal distribution. The density functions and normal Q-Q-plots[23] in Appendix 3.C show clear departure of the residuals from the Gaussian norm, especially for Model 1 and especially in the tails. Non-normalities are of lesser extent for Model 3, but we still observe extreme realizations in the upper quantiles. This motivates our interest to

[22] We experimented with different quantile orders, i.e. $\tau = 0.1, 0.25, 0.5, 0.75, 0.9$ and $\tau = 0.1, 0.2, 0.4, 0.6, 0.8$. The maximum quantile order is however restricted to 10 given the limited sample size.

[23] Normal Q-Q-plots compare the sample quantiles on the vertical axis to the theoretical quantiles of the standard normal distribution on the horizontal axis

further explore the conditional distribution of inflation. In the case of non-normal errors, non-linear estimators might be more efficient estimators than OLS.

Quantile regression estimates for Models 1 and 3 are reported in Appendix 3.D. Slope equality tests based on Koenker and Bassett (1982) reveal that coefficients differ across quantiles, especially for Model 1.[24]

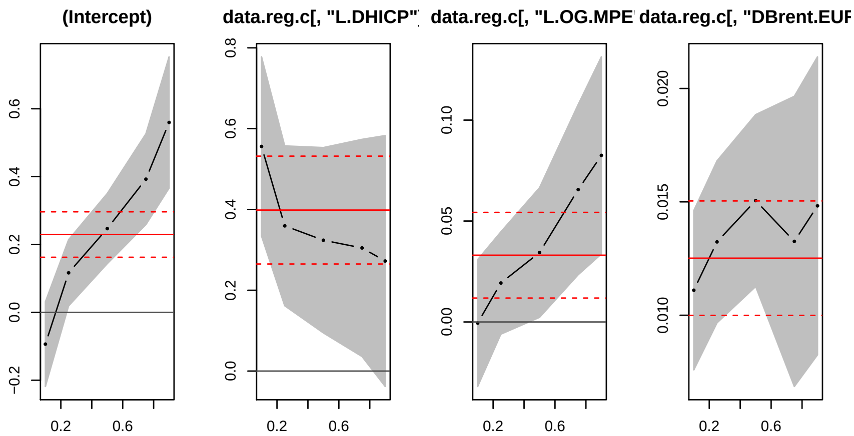


Figure 3.4: Quantile slopes for the dependent variables for Model 1 for headline inflation

Quantile regression estimates are summarized into graphical representations of quantile-specific slopes in Figure 3.4 (see Appendix 3.D for the quantile slopes of the other specifications). We see important variations in the impact of the dependent variables over different quantiles, even if they are not always significantly different from the mean impact. We observe notably a - significantly - higher impact of lagged inflation on current inflation at the left tail of the distribution. This implies that inflation is more persistent during low inflation times. This is in line with the findings in Busetti et al. (2015). Regarding the output gap, we find an increasing response of inflation to (lagged) domestic slack at the right tail of the distribution, i.e. domestic slack measures seem to play a stronger role during high inflation times. The (contemporaneous) impact of the global factor on inflation varies somewhat less over the conditional inflation quantiles. For the model with oil prices (Model 1), the impact increases towards the median, before falling back again at higher quantiles. For the model with import prices (Model 3), the impact is relatively stable and close to the median across the conditional distribution before jumping up abruptly at the very high end of the distribution (i.e. for high levels of inflation). This diverging impact of the dependant variables on inflation across conditional quantiles implies that when assuming mean coefficients, we may either over- or underestimate their impact on inflation. If inflation was located at the lower end of the conditional distribution, we might notably underestimate the impact of past inflation. Also, during

[24]This result is somewhat sensitive to the chosen quantile order. The Chi-square test rejects the null hypothesis of equal quantile coefficients at the 5%-level for quantile order $\tau = 0.1, 0.2, 0.4, 0.6, 0.8$ but not for quantile order $\tau = 0.1, 0.25, 0.5, 0.75, 0.9$ for Model 1. But even for the latter, pairwise comparison tests across quantiles still reveal significant differences in the quantile coefficients for some of the quantiles.

low inflation periods, the impact of domestic slack and global factors might be somewhat smaller than during normal and high inflation times. The higher persistency of the inflation process at the left tail of the distribution could also explain why Ciccarelli and Osbat (2017) find an increase of the persistency of inflation over the most recent period of low inflation from 2014 to 2016.

3.5.3 Forecast performance

Quantile regressions are generally used to derive forecast distributions based on the forecast quantiles. We consider an alternative approach to evaluate whether the quantile regression results can be useful for forecasting purposes. We conduct point forecasts from quantile regressions based on specific quantiles and compare it to the forecasts from OLS.[25] We concentrate our analysis on the most recent period of low inflation from 2014 to 2016 and use the quantile regression model corresponding to the quantile of past observed inflation. Hence, to forecast inflation for $t+1$ with quantile regressions, we rely on the model corresponding to the quantile of observed inflation at date t, based on the historical distribution of inflation up to date t. For instance, in 2013Q4, headline inflation was located at the extreme left of the distribution (first quantile). We then conduct the one-quarter-ahead forecast for 2014Q1 based on the model corresponding to the first quantile and compare it to observed inflation. We then roll the estimation one quarter forward, recalculate the historical distribution up to 2014Q1 and make forecasts based on the model for the observed inflation quantile in 2014Q1, and so forth. We analyse two forecast periods: (i) the period from 2014 to 2015, where inflation was persistently low and located most of the time in the first quantile of the distribution (except for the one quarter in 2015Q3), and (ii) the period from 2014 to 2016, where inflation rebounded.

	Dependent variable	Headline HICP	
	Forecast period	2014-2015	2014-2016
	Model		
PC-OG	OG	0.78	1.17
M1	OG+oil price	0.74	1.11
M3	OG+import prices	0.77	1.17
M5	OG+OECD CPI ex. EA (weight.)	0.68	1.01
M7	OG+FDR	0.74	1.17

Models are estimated on rolling windows of a fixed size of 71 quarters. Mean absolute forecast errors (MAE) for $h = 1$ are computed over two forecast periods: 2014Q1 to 2015Q4 (8 observations) and 2014Q1-2016Q4 (12 observations). Ratios below 1 signify a lower MAE for the quantile regressions compared to OLS. OG stands for output gap.

Table 3.5: MAE ratios between quantile regressions and OLS for headline HICP ($h = 1$)

[25] We focus on the one-quarter-ahead forecast for which the quantile models are better identified.

Table 3.5 reports the relative mean-absolute forecast errors (MAE) for quantile regressions compared to the corresponding OLS estimates. For the period of persistently low inflation from 2014 to 2015, forecasts produced by quantile regressions are superior to the corresponding OLS forecasts for all of the specifications. The gains in forecast accuracy are quite large, with improvements in the MAE of between 22% and 32%. The best overall forecast model for this period (Quantile Model 7 with the foreign demand index) has a MAE of 0.20 compared to 0.27 for the corresponding OLS estimate with Model 7 (the best OLS model) and to 0.30 for the OLS estimate of the standard Phillips curve. We also find that most of the quantile models augmented with global factors (Model 1, 5 and 7) outperform the quantile model including only the domestic slack variable. These results change however when we look at the longer forecast period from 2014 to 2016, which includes the pick-up of inflation in 2016. For this period, quantile regressions no longer outperform the forecast ability of the corresponding OLS estimates. The lowest forecast error of 0.26 is achieved with the OLS Model 7, compared to a MAE of 0.29 for the quantile model with the lowest MAE (Model 5) and of 0.30 for the quantile model 7. These results leave us to conclude that quantile regressions can be a useful addition to OLS for forecasting short-term headline inflation in periods of persistently low (or high) inflation. Quantile regressions are of lesser use for forecasting during periods of volatile inflation. In such periods, quantile regressions do not provide better forecast results than OLS.

3.6 Conclusions

In this chapter, we examined the fit and the forecast ability of the Phillips curve augmented with different global factors for euro area inflation in a systematic manner, focusing on the period before the COVID crisis and the recent inflation surge. We show that global factors generally improve the fit of the Phillips curve for headline inflation but that they are less relevant for forecasting, with one possible exception of trade-weighted foreign demand. While introducing traditional global indicators such as commodity prices and import prices or global consumer inflation into the Phillips curve improves its fit, we find little support for introducing global slack measures into the Phillips curve, once domestic slack and more direct measures of global price pressures have been taken into account. Regarding core inflation, the importance of global factors is considerably reduced and difficult to establish in a significant manner in a Phillips curve framework. Anticipation measures from surveys or forecasts contain more useful information for explaining core inflation. Overall, we conclude that the Phillips curve remains a useful tool for studying and forecasting inflation. The forecast behaviour of the Phillips curve in its standard or augmented form is however not stable over time, especially as regards headline inflation, and it performs better during some periods than during others. Turning to quantile considerations, our results confirm the interest of exploring the entire conditional distribution of inflation in addition to standard conditional mean estimation by OLS. We show that the tail behaviour of some of the predictors such as the lagged inflation term and domestic slack do matter and that there exists asymmetries in the reaction of inflation to these predictors. Forecasts from quantile regressions can be a useful addition to forecasts from OLS in some periods. This result could be exploited further by extending the analysis to different time periods or contexts. Further alleys to investigate concern notably a more exhaustive description of the macroeconomic environment (crisis vs. normal times, high vs. low dynamics of global forces, demand vs. supply shocks), in addition to the inflationary regime, which determines the forecast ability of the Phillips curve. The recent rise in inflation is a natural candidate to test some premises about the empirical performance of the augmented linear and non-linear Phillips curve.

Appendix

3.A Data sources and descriptions

Description	Source	Details
	Euro area inflation	
HICP headline	Eurostat	sa(X12)
HICP ex. energy and food	Eurostat	sa(X12)
HICP ex. energy	Eurostat	sa(X12)
Household's inflation expectations	European Commission	sa, price trends over next 12 months
Consensus forecast	Consensus Economics	One quarter and four quarters ahead forecast
	Euro area slack	
EA Unemployment rate	Eurostat	sa
EA unemployment gap	ECB & authors' calc.	Difference between NAIRU and unemployment rate
EA OG	European Commission	Cubic spline
EA OG	ECB	Cubic spline
EA OG	Eurostat & authors' calc.	Real GDP sa, HP filtered
EA industrial production	Eurostat	sa
	Global factors	
	Commodities	
Oil price, USD	ECB	Brent crude oil price in USD
Oil price, EUR	ECB	Brent crude oil price in EUR
Price of non-energy commodities, EUR	ECB	
	Exchange rates	
USD/EUR exchange rate	ECB	Exchange rate against euro, spot (mid)
Nominal effective exchange rate	ECB	Exchange rate against 12 trading partners
Nominal effective exchange rate	ECB	Exchange rate against 19 trading partners
Nominal effective exchange rate	ECB	Exchange rate against 38 trading partners
	Foreign demand and global slack	
Foreign demand index (FDR)	ECB	Trade-weighted imports of EA trading partners
OG, US	IMF	Cubic spline
OG, advanced economies*	IMF & authors' calc.	GDP-weighted, cubic spline
OG, adv. economies ex. EA	IMF & authors' calc.	GDP-weighted, cubic spline
OG, emerging markets	IMF & authors' calc.	GDP-weighted, cubic spline
OG, US-JPN-UK-CAN	IMF & authors' calc.	GDP-weighted, cubic spline
OG, world	IMF	Cubic spline
OG, world ex. EA	IMF & authors' calc.	GDP-weighted, cubic spline
US unemployment rate	OECD	sa
OECD unemployment rate	OECD	Unemployment rate of advanced economies, sa
Unemployment rate, OECD ex. EA	OECD & authors' calc.	sa, GDP-weighted
Unemployment rate, US-JPN-UK-CAN	OECD & authors' calc.	sa, GDP-weighted
	Global inflation	
OECD CPI ex. EA	OECD & authors' calc.	Simple or weight. avg., exp.-weighted, sa(X12)
CPI, US-JPN-UK-CAN	OECD & authors' calc.	Simple or weight. avg., exp.-weighted, sa(X12)
US CPI	OECD & authors' calc.	sa(X12)
World** CPI ex. EA	OECD & authors' calc.	Simple or weight. avg., exp.-weighted, sa(X12)
OECD CPI (ex. EA) ex. food & energy	OECD & authors' calc.	Simple or weight. avg., exp.-weighted, sa(X12)
World CPI (ex. EA) ex. food & energy	OECD & authors' calc.	Simple or weight. avg., exp.-weighted, sa(X12)
CPI ex. food & energy, US-JPN-UK-CAN	OECD & authors' calc.	Simple or weight. avg., exp.-weighted, sa(X12)
US CPI ex. food & energy	OECD & authors' calc.	sa(X12)
Competitors' prices, USD	ECB	Export prices of extra EA trading partners in USD
Competitors' prices, EUR	ECB	Export prices of extra EA trading partners in EUR
Import deflator, EUR	Eurostat	Extra EA import deflator
Relative import deflator, EUR	Eurostat & authors' calc.	Extra EA Import deflator against GDP deflator

*Australia, Canada, Denmark, Japan, Korea, New Zealand, Norway, Sweden, the United Kingdom, the United States and the euro area.
**Members of the G20 and members of the OECD (Australia, Brazil, Canada, Chile, China, Colombia, Denmark, Hungary, Iceland, India, Indonesia, Israel, Japan, Korea, Lithuania, Mexico, New Zealand, Norway, Poland, Russia, Saudi Arabia, South Africa, Sweden, Switzerland, Turkey, the United Kingdom, the United States and EA).
Note: CPI stands for Consumer price index, HICP for Harmonized index of consumer prices, OG for output gap, sa for seasonally adjusted

3.B OLS estimates of augmented Phillips curves

Headline inflation

Models	AR1	PC(OG)	PC(UR)	M1	M2	M3	M4	M5	M6	M7	M8
Intercept	0.22***	0.27***	0.89**	0.24***	0.28***	0.31***	-0.06	0.04	0.13***	0.24***	0.38***
Lagged inflation	0.45***	0.36**	0.34***	0.38***	0.30**	0.20**	0.20*	0.18**	0.20*	0.20**	0.34***
Lagged EA OG / UR		0.04*	-0.06*	0.03*	0.04*	0.05***	0.02	0.04**	0.03*	0.06***	0.03
Oil price EUR				0.01***							
Non-energy commodity prices					0.02***						
Import prices						0.08***					
OECD CPI ex. EA (simple)							0.55***				
OECD CPI ex. EA (weight.)								0.54***			
US CPI									0.38***		
FDR										0.07***	
Lagged OG adv. ex. EA											0.00
Adjusted R2	0.20	0.23	0.24	0.60	0.37	0.56	0.49	0.59	0.60	0.41	0.23
Observations	82	82	82	82	82	82	82	82	82	82	82

Models	M9	M10	M11	M12	M13	M14
Intercept	0.30	0.17	0.25**	0.12*	1.03***	0.09
Lagged inflation	0.35**	0.35**	0.33**	-0.36***	0.22**	-0.37***
Lagged EA OG / UR	0.04	0.03*	0.04	0.00	-0.08***	-0.02
OCDE CORE CPI ex. EA	-0.06					
US CORE CPI		0.18				
Consumer infl. expectations			0.00			
Consensus				0.41***		0.39***
Foreign demand					0.07***	
Adjusted R2	0.23	0.23	0.23	0.55	0.40	0.55
Observations	82	82	82	82	82	82

Notes: Stars *, ** and *** denote statistical significance at 10%, 5% and 1% levels respectively. Models are estimated over the 1996Q3-2016Q4 sample. Commodity prices and import prices are in EUR. Variables are in log-difference, except for the output gap, which is introduced in levels.

Table 3.B.1: Estimation results for Phillips curves for headline infation

Dependent variable: HICP excluding energy (HEX)

Models	AR1	PC(OG)	PC(UR)	M1	M2	M3	M4	M5	M6	M7	M8
Intercept	0.14***	0.23***	0.64***	0.22***	0.23***	0.21***	0.16**	0.19***	0.20***	0.21***	0.23***
Lagged inflation	0.63***	0.10***	0.45***	0.41***	0.40***	0.43***	0.39***	0.40***	0.41***	0.40***	0.40***
Lagged EA OG / UR		0.03**	-0.04***	0.03**	0.03**	0.03**	0.03**	0.03**	0.03**	0.03***	0.03*
Oil price EUR				0.00							
Non-energy commodity prices					0.00						
Import prices						0.01					
OECD CPI ex. EA (simple)							0.10*				
OECD CPI ex. EA (weight.)								0.07*			
US CPI									0.04*		
FDR										0.01	
Lagged OG adv. ex. EA											0.00
Adjusted R2	0.38	0.46	0.45	0.45	0.46	0.48	0.49	0.47	0.47	0.46	0.46
Observations	82	82	82	82	82	82	82	82	82	82	82

Models	M9	M10	M11	M12	M13	M14
Intercept	0.23***	0.20**	0.20***	0.15***	0.63***	0.20**
Lagged inflation	0.40***	0.40***	0.25*	0.20*	0.45***	0.40***
Lagged EA OG / UR	0.03**	0.03**	0.03***	0.02**	-0.04***	-0.03**
OCDE CORE CPI ex. EA	-0.01					
US CORE CPI		0.05				
Consumer infl. expectations			0.01**			
Consensus				0.09***		0.05***
Foreign demand					0.01	
Adjusted R2	0.45	0.45	0.53	0.55	0.45	0.45
Observations	82	82	82	82	82	82

Notes: Stars *, ** and *** denote statistical significance at 10%, 5% and 1% levels respectively. Models are estimated over the 1996Q3-2016Q4 sample. Commodity prices and import prices are in EUR. Variables are in log-difference, except for the output gap, which is introduced in levels.

Table 3.B.2: Estimation results for Phillips curves augmented for inflation excluding energy

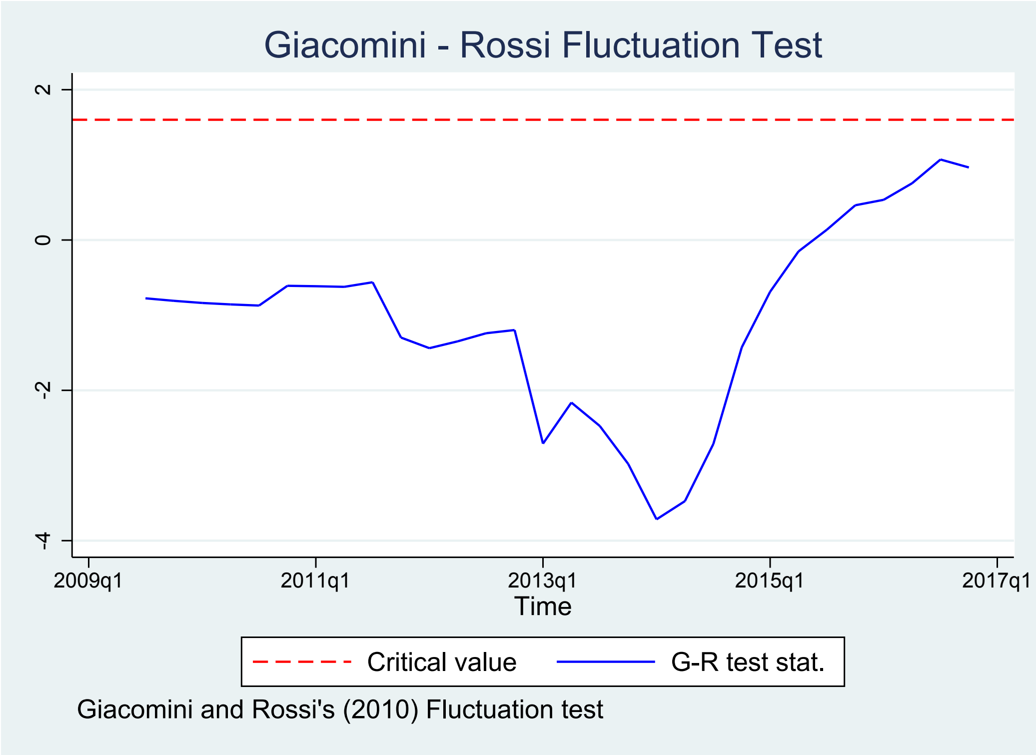
Dependent variable: HICP excluding energy and food CORE)

Models	AR1	PC(OG)	PC(UR)	M1	M2	M3	M4	M5	M6	M7	M8
Intercept	0.14***	0.24***	0.60***	0.24***	0.24***	0.23***	0.20***	0.23***	0.21***	0.23***	0.24***
Lagged inflation	0.58***	0.31***	0.37***	0.30***	0.31***	0.32***	0.32***	0.31***	0.31***	0.31***	0.31***
Lagged EA OG / UR		0.03***	-0.04***	0.03**	0.03***	0.03***	0.03**	0.03***	0.03**	0.03***	0.03***
Oil price EUR				0.00							
Non-energy commodity prices					0.00						
Import prices						0.00					
OECD CPI ex. EA (simple)							0.05				
OECD CPI ex. EA (weight.)								0.02			
US CPI									0.01		
FDR										0.00	
Lagged OG adv. ex. EA											0.00
Adjusted R2	0.33	0.44	0.43	0.43	0.43	0.43	0.44	0.43	0.43	0.44	0.43
Observations	82	82	82	82	82	82	82	82	82	82	82

Models	M9	M10	M11	M12	M13	M14
Intercept	0.23***	0.21***	0.20***	0.19***	0.60***	0.47***
Lagged inflation	0.31***	0.31***	0.29**	0.25*	0.37***	0.27***
Lagged EA OG / UR	0.03***	0.03***	0.02***	0.02***	-0.04***	-0.03***
OCDE CORE CPI ex. EA	0.00					
US CORE CPI		0.04				
Consumer infl. expectations			0.00**			
Consensus				0.04**		0.04***
Foreign demand					0.00	
Adjusted R2	0.43	0.43	0.47	0.47	0.42	0.47
Observations	82	82	82	82	82	82

Notes: Stars *, ** and *** denote statistical significance at 10%, 5% and 1% levels respectively. Models are estimated over the 1996Q3-2016Q4 sample. Commodity prices and import prices are in EUR. Variables are in log-difference, except for the output gap, which is introduced in levels.

Table 3.B.3: Estimation results for Phillips curves for inflation excluding energy and food

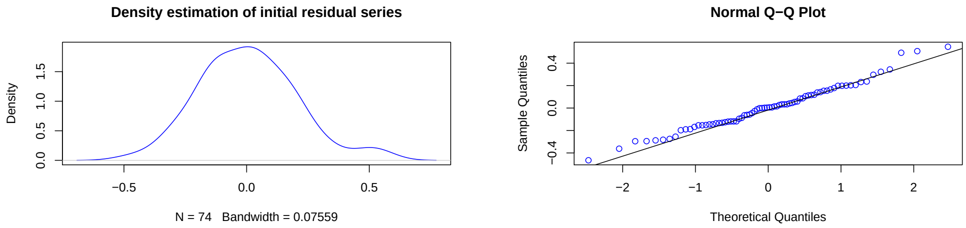


Note: Giacomini and Rossi's one-sided fluctuation test for the standard Phillips curve and the AR over 15-quarters rolling forecast window for $h = 1$. The test rejects the null hypothesis of equal predictive ability when the test statistic is above the band line.

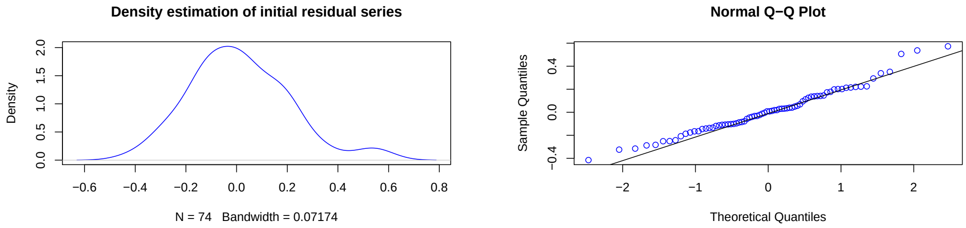
Figure 3.B.1: Giacomini and Rossi's fluctuation test

3.C Quantile densities

Density plots of residuals from estimated model M1 and variations

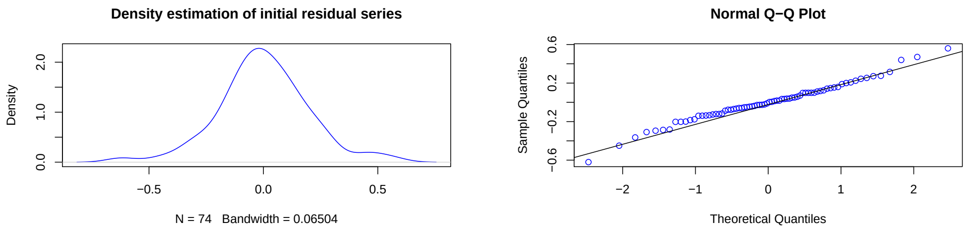


Model 1 with OG

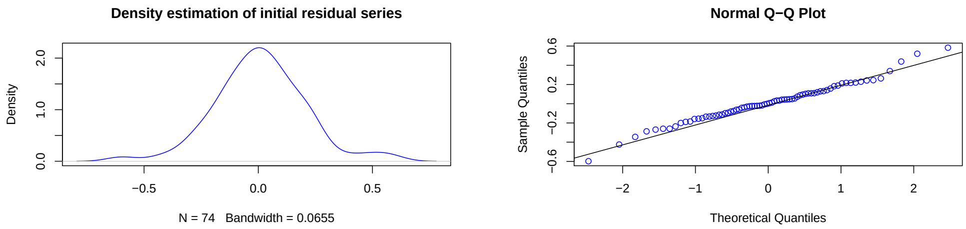


Model 1 with UR

Density plots of residuals from estimated model M3 and variations



Model 3 with OG



Model 3 with UR

Figure 3.C.1: Density plots and normal Q-Q-plots for different models

3.D Quantile inference

Table 3.D.1: Quantile regression estimates for headline inflation

	Dependent variable: Headline inflation				
Quantile τ	0.10	0.25	0.50	0.75	0.90
			PC1		
Intercept	-0.113 (0.108)	0.087 (0.093)	0.243 (0.089)	0.400 (0.098)	0.701(0.141)
Lagged inflation	0.516 (0.207)	0.500 (0.202)	0.448 (0.176)	0.362 (0.193)	0.059 (0.255)
Lagged EA Output gap	0.004 (0.041)	0.041 (0.026)	0.038 (0.024)	0.047 (0.027)	0.064 (0.038)
Pseudo R2	0.15	0.18	0.17	0.15	0.16
			Model 1 Oil price in EUR		
Intercept	-0.076 (0.077)	0.117 (0.059)	0.242 (0.082)	0.425 (0.072)	0.557 (0.103)
Lagged inflation	0.531 (0.128)	0.366 (0.124)	0.337 (0.163)	0.247 (0.142)	0.265 (0.170)
Lagged EA Output gap	-0.002 (0.022)	0.018 (0.016)	0.032 (0.024)	0.058 (0.024)	0.063 (0.031)
Oil price EUR	0.011 (0.002)	0.013 (0.002)	0.014 (0.002)	0.013 (0.003)	0.013 (0.003)
Pseudo R2	0.52	0.45	0.38	0.32	0.32
			Model 3 Import prices		
Intercept	0.075 (0.087)	0.161 (0.066)	0.322 (0.070)	0.436 (0.075)	0.589 (0.098)
Lagged inflation	0.224 (0.158)	0.297 (0.142)	0.211 (0.143)	0.179 (0.162)	0.105 (0.186)
Lagged EA Output gap	0.043 (0.031)	0.041 (0.015)	0.055 (0.019)	0.057 (0.025)	0.069 (0.033)
EA import prices	0.085 (0.020)	0.082 (0.024)	0.074 (0.023)	0.088 (0.021)	0.093 (0.025)
Pseudo R2	0.39	0.38	0.35	0.32	0.32
			Model 5 OECD CPI ex.EA (weighted average)		
Intercept	-0.151 (0.094)	-0.088 (0.062)	-0.007 (0.065)	0.176 (0.073)	0.264 (0.123)
Lagged inflation	0.191 (0.174)	0.277 (0.128)	0.268 (0.137)	0.091 (0.140)	0.059 (0.179)
Lagged EA Output gap	0.025 (0.023)	0.027 (0.017)	0.030 (0.019)	0.028 (0.021)	0.034 (0.026)
OECD inflation ex. EA	0.511 (0.117)	0.500 (0.092)	0.575 (0.097)	0.611 (0.129)	0.623 (0.185)
Pseudo R2	0.42	0.41	0.37	0.33	0.33
			Model 7 Foreign demand index		
Intercept	-0.111 (0.010)	0.007 (0.072)	0.181(0.052)	0.334 (0.117)	0.687 (0.141)
Lagged inflation	0.277 (0.200)	0.304 (0.133)	0.374 (0.117)	0.349 (0.165)	0.250 (0.187)
Lagged EA Output gap	0.043 (0.032)	0.046 (0.018)	0.044 (0.016)	0.046 (0.023)	0.055 (0.040)
FDR	0.099 (0.025)	0.108 (0.022)	0.057 (0.015)	0.031 (0.027)	-0.046 (0.053)
Pseudo R2	0.38	0.32	0.28	0.18	0.14

Notes: Estimated coefficients for quantile orders $\tau = 0.10, 0.25, 0.50, 0.75, 0.90$. Estimation period: 1996-2016. Standard errors computed by bootstrap are reported in brackets.

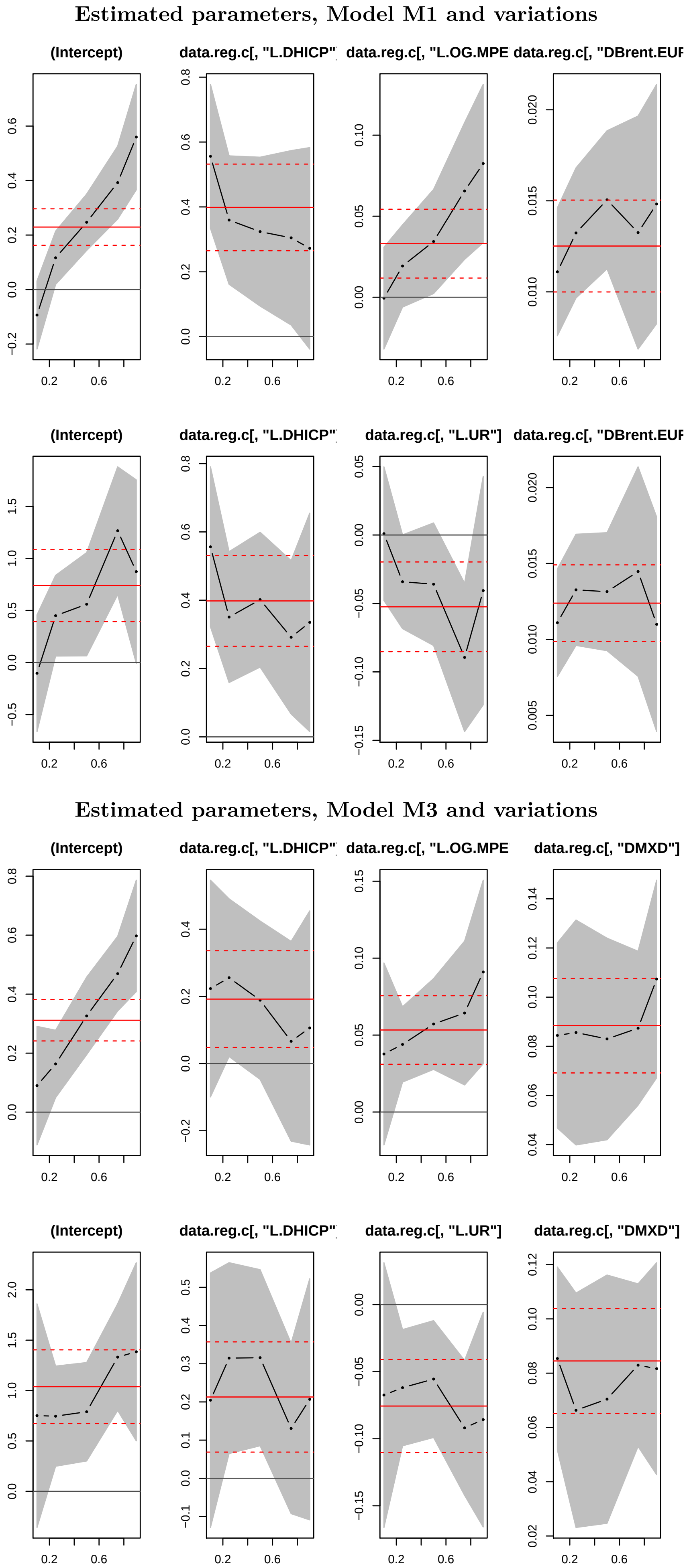


Figure 3.D.1: Quantile slopes for different models

General Conclusion

In this thesis, I have studied how national fiscal policies, fiscal spillovers and global factors influence the stabilisation of output and prices in the entire euro area and in individual euro area member countries. I have shown that expenditure-driven national fiscal policies have non-negligible output and price spillovers in the euro area, affecting the fiscal stance and fiscal stabilisation properties in other member countries. These spillovers are greater the more integrated and open member countries, from which the fiscal impulse originates, are. The instrument of fiscal policy also matters: government consumption generates slightly higher spillovers than government investment. The real effective exchange rate is found to adjust only slowly to national fiscal policies in a currency union, considering that nominal exchange rates are fixed and relative prices are sticky. I have also shown that discretionary fiscal policies have failed, on average, to deliver counter-cyclical macroeconomic stabilisation in the euro area. Fiscal outcomes were broadly a-cyclical since the start of EMU. This has increased the potentiel burden on monetary policy for smoothing the cycle in the euro area. Budget plans were even found to be pro-cyclical on average, and this also when excluding the European Debt crisis from the analysis, where pro-cyclical consolidation is generally understood as a cause for the deepening of the crisis (Bartsch et al., 2020). Pro-cyclicality is a consequence of market stress and signalling effects. However, the result of pro-cyclicality is not uniform across euro area countries, and I have shown that some member countries perform better than others in setting up counter-cyclical fiscal plans. Blanchard et al. (2021) suggest that this, on average, ineffective macroeconomic stabilisation by fiscal policy is already inherent in the institutional setup of the euro area because countries do not take sufficiently into account spillovers of their fiscal action to other member countries and to the euro area aggregate.[26] However, our finding that pro-cyclicality is already anchored in budget plans speaks against the role of spillovers as an explanation for the lack of stronger counter-cyclical stabilisation, considering that countries' budget plans exert only a limited influence on other countries' budget plans. In other words, actual output spillovers do not drive the result of pro-cyclicality. Furthermore, I have illustrated that global factors influence the aggregate price level in the euro area and that these factors can best be represented by commodity prices and import prices in a Phillips curve framework. Broader indictors such as global

[26] Indeed, the only policy coordination mechanism, which exists in the euro area, is the European Semester. It has however been described as largely ineffective (Darvas and Leandro, 2015).

slack measures or global consumer prices do not improve the understanding of euro area inflation dynamics. I have also demonstrated that the inflation process is not linear in the euro area. Price pressures from domestic real activity and partly also from global factors become greater when inflation is higher. With regard to the last point, a consensus is slowly emerging in the literature on the non-linearity of the Phillips curve (see recently, Harding et al., 2023 and Dao et al., 2023).

The work presented here might be extended and complemented in several ways.

First, the analyses in chapter one dealing with macroeconomic stabilisation could usefully be supplemented by an empirical investigation of interactions of monetary and fiscal policy in the euro area. This analysis of monetary and fiscal interactions has recently gained more attention in the literature. Kloosterman et al. (2022), for example, assess empirically how the effects of monetary easing and tightening are affected by the prevailing fiscal stance in the euro area. They find that the responses of inflation and output growth to a monetary policy shock depend on whether the economy is in a contractionary or expansionary fiscal regime. Reichlin et al. (2023) study the fiscal response to both conventional and unconventional monetary policy in the euro area and their consequent effects on output and prices in the 4 main euro area economies. Afonso et al. (2019) study interactions between monetary and fiscal policies in the EU by including monetary and fiscal variables in the other's authority reaction function. ECB (2021) provide a more general discussion and model simulations for the monetary-fiscal mix in the euro area. Further work could study this issue in real time, i.e. at the time of decision making vs. outcomes, using the real time dataset presented in chapter one. In an ongoing work, I look at this question and analyse the fiscal response to monetary conditions, both for conventional and unconventional monetary policy indicators (Schmidt, 2023). I also allow for a non-linear reaction of fiscal policy to monetary conditions using a panel threshold model. My preliminary findings show that fiscal policy complemented conventional monetary policy in the euro area in macroeconomic stabilisation, i.e. it moved in the same direction. Fiscal authorities tightened (respectively, eased) the fiscal stance in response to an increase (respectively, decrease) of monetary policy rates on average since the start of EMU. No significant fiscal reaction is put into evidence for unconventional monetary policy indicators. I also do not detect any threshold effects in the fiscal response to monetary policy conditions.

Second, and as mentioned above, potential causes for the empirical outcome of pro-cyclical policies in developed countries, and specifically in the euro area, could be further investigated. Gootjes and de Haan (2023) provide a recent work in this regard. Another option would be to better analyse the cyclicality of the two sides of the fiscal balance, i.e. expenditures and revenues, separately. In other words, is pro-cyclicality due to an over- or under-adjustment of discretionary expenditures, or is it the result of a pro-cyclical reaction of discretionary revenues? A narrative approach, instead of the more aggregate CAB approach used here, could also be employed to calibrate the size of the discretionary fiscal shock[27] and better capture causes of pro-cyclicality. The literature has also brought forward different political and institutional variables to explain the cyclicality of fiscal policy, which could be extended further (see for a literature review and a recent analysis of the role of EU fiscal rules Larch et al., 2021b). The role of implementation lags, for example, has so far only been insufficiently analysed in the empirical literature, especially with a perspective on developed countries. More specifically in the euro area, the institutional setup, and in particular the role of coordination or coordination failure, could also be further explored.

Third, in terms of the analysis of fiscal spillovers, the interaction with the monetary policy response is an important question, given the implications of spillovers for prices. Blagrave et al. (2017) have provided some insights in this regard and have shown that fiscal spillovers are larger when monetary policy is close to the effective lower bound. There is room to expand this analysis. Also, the different size of spillovers found with empirical models compared to structural models is a question of interest. In fact, studies based on DSGE models often found spillovers to be much lower than VAR-based estimates. I have addressed this question with my coauthors in a joint work, where we use different modelling approaches, of which a DSGE model with an extended fiscal block calibrated to euro area countries, to estimate spillovers in the euro area (Alloza et al., 2020). The model simulations have shown that the difference in the size of spillovers between models can be largely explained by the reaction of interest rates. Spillovers are higher in a DSGE model when monetary policy is deactivated, i.e. interest rates do not react.

Fourth, concerning the third chapter, updating the work with data for the latest high inflation period in the euro area would be interesting. Fifth, studying the impact of global influences, especially global shocks, on euro area inflation remains highly relevant. In this sense, a richer VAR modelling approach, along the lines of Bianchi and Civelli (2015), could be used. The underlying global indicators could also be further refined to give an even better insight into the transmission of global price pressures to domestic inflation. In particular, indicators related to global supply chain integration received an increasing attention in the literature in recent years (see, for example, ECB, 2019 and Andrews et al., 2018). There have become even more important during the recent period of high inflation, given the significant role of global supply chain disruptions in the rise in inflation. Finally, the identification of the Phillips curve remains an important area of research, for the euro area and in general, given its wide use in the ECB and other central banks as forecasting and policy tool (see "General Introduction" for some research innovations on the identification of the Phillips curve).

In general, the interaction between monetary and fiscal policy remains a key question for the euro area in the face of new global challenges and shocks and given its specific institutional setup.

[27] Narrative measures are more difficult to obtain, especially for a panel of countries, which complicates the use of this method.

Bibliography

Afonso, A., Alves, J., and Balhote, R. Interactions between monetary and fiscal policies. *Journal of Applied Economics*, 22(1):132–151, 2019.

Aldama, P. and Creel, J. Real-time fiscal policy responses in the OECD from 1997 to 2018: Procyclical but sustainable? *European Journal of Political Economy*, 73(102135), 2022.

Alesina, A. and Tabellini, G. Why is fiscal policy often procyclical? NBER Working Papers 11600, National Bureau of Economic Research, Inc., 2005.

Alloza, M., Buriel, P., and Perez, J. J. Fiscal policies in the euro area: Revisiting the size of spillovers. *Journal of Macroeconomics*, 61(103132), 2019.

Alloza, M., Ferdinandusse, M., Jacquinot, P., and Schmidt, K. Fiscal expenditure spillovers in the euro area: An empirical and model-based assessment. Occasional Paper Series 240, European Central Bank, April 2020.

Andrews, D., Gal, P., and Witheridge, W. A genie in a bottle?: Globalisation, competition and inflation. OECD Economics Department Working Papers 1462, OECD Publishing, March 2018.

Auer, R., Borio, C., and Filardo, A. The globalisation of inflation: The growing importance of global value chains. BIS Working Papers 602, Bank for International Settlements, 2017.

Auerbach, A. J. and Gorodnichenko, Y. Measuring the output responses to fiscal policy. *American Economic Journal: Economic Policy*, 4(2):1–27, 2012.

Auerbach, A. J. and Gorodnichenko, Y. Output spillovers from fiscal policy. *American Economic Review*, 103(3):141–146, 2013.

Ball, L. M. and Mazumder, S. A Phillips curve for the euro area. NBER Working Paper 26450, National Bureau of Economic Research, Inc., 2019.

Banbura, M. and Bobeica, E. Does the Phillips curve help to forecast euro area inflation? *International Journal of Forecasting*, 39(1):364–390, 2023.

Banbura, M. and Mirza, H. Forecasting euro area inflation with the Phillips curve. Mimeo, 2013.

Bankowski, K. and Ferdinandusse, M. Euro area fiscal stance. Occasional Paper Series 182, European Central Bank, 2017.

Barbier-Gauchard, A. and Betti, T. Spillover effects of fiscal policy in a monetary union: Why do fiscal instruments matter? *Bulletin of Economic Research*, 73(1):1–33, 2020.

Bartsch, E., BénassyQuéré, A., Corsetti, G., and Debrun, X. It's all in the mix: How monetary and fiscal policies can work or fail together. Geneva Report 23, CEPR Press, 2020.

Beetsma, R. and Giuliodori, M. Fiscal adjustment to cyclical developments in the OECD: An empirical analysis based on real-time data. *Oxford Economic Papers*, 62(1):419–441, 2010.

Beetsma, R. and Giuliodori, M. The effects of government purchases shocks: Review and estimates for the EU. *The Economic Journal*, 121(550):4–32, 2011.

Beetsma, R., Giuliodori, M., and Klaasesen, F. Trade spillovers of fiscal policy in the European Union: A panel analysis. *Economic Policy*, 21(48):639–687, 2006.

Berger, H. and Nitsch, V. Zooming out: The trade effect of the euro in historical perspective. *Journal of International Money and Finance*, 27(8):1244–1260, December 2008.

Bernoth, K., Hughes, A., and Lewis, J. Did fiscal policy makers know what they were doing? Reassessing fiscal policy with real time data. CEPR Discussion Paper 6758, CEPR, 2008.

Bianchi, F. and Civelli, A. Globalization and inflation: Evidence from a time varying VAR. *Review of Economic Dynamics*, 18(2):406–433, 2015.

Blagrave, P., Ho, G., Koloskova, K., and Vesperoni, E. Fiscal spillovers: The importance of macroeconomic and policy conditions in transmission. Spillover Note 11, International Monetary Fund, 2017.

Blanchard, O. and Perotti, R. An empirical characterization of the dynamic effects of changes in government spending and taxes on output. *The Quarterly Journal of Economics*, 117(4):1329–1368, 2002.

Blanchard, O., Leandro, A., and Zettelmeyer, J. Redesigning EU fiscal rules: From rules to standards. *Economic Policy*, 36(106):195–236, 2021.

Bobeica, E. and Hartwig, B. The COVID-19 shock and challenges for time series models. Working Paper Series 2558, European Central Bank, 2021.

Bonam, D., Haan, J. D., and Soederhuizen, B. The effects of fiscal policy at the effective lower bound. *Macroeconomic Dynamics*, 26(1):149–185, January 2022.

Borio, C. and Filardo, A. Globalisation and inflation: New cross-country evidence on the global determinants of domestic inflation. BIS Working Papers 227, Bank for International Settlements, 2007.

Breusch, T. and Pagan, A. R. The lagrange multiplier test and its applications to model specification in econometrics. *Review of Economic Studies*, 47(1):239–253, 1980.

Bun, M. J. G. and Kiviet, J. F. The effects of dynamic feedbacks on LS and MM estimator accuracy in panel data models. *Journal of Econometrics*, 132(2):409–444, 2006.

Busetti, F., Caivano, M., and Rodano, L. On the conditional distribution of euro area inflation forecast. Temi di discussione (Working papers) 1027, Bank of Italy, 2015.

Busetti, F., Caivano, M., and Delle-Monache, D. Domestic and global determinants of inflation: Evidence from expectile regression. *Oxford Bulletin of Economics and Statistics*, 83(4):982–1001, 2021.

Canova, F. and Ciccarelli, M. Panel vector autoregressive models: A survey. Working Paper Series 1507, European Central Bank, 2013.

Cassette, A., Creel, J., Farvaque, E., and Paty, S. Governments under influence: Country interactions in discretionary fiscal policy. *Economic Modelling*, 30:79–89, 2013.

Chahad, M., Hofmann-Drahonsky, A.-C., Meunier, B., Page, A., and Tirpák, M. What explains recent errors in the inflation projections of Eurosystem and ECB staff? Economic Bulletin Articles 3, European Central Bank, 2022.

Chatelais, N. and Schmidt, K. The impact of import prices on inflation in the euro area. Rue de la Banque 37, Banque de France, 2017.

Checherita-Westphal, C. and Zdarek, V. Fiscal reaction function and fiscal fatigue: Evidence for the euro area. Working Paper Series 2036, European Central Bank, 2017.

Ciccarelli, M. and Mojon, B. Global inflation. *The Review of Economics and Statistics*, 92(3):524–535, 2010.

Ciccarelli, M. and Osbat, C. Low inflation in the euro area: Causes and consequences. Occasional Paper Series 181, European Central Bank, 2017.

Cimadomo, J. Fiscal policy in real time. *The Scandinavian Journal of Economics*, 114 (2):440–465, 2012.

Cimadomo, J. Real-time data and fiscal policy analysis: A survey of the literature. *Journal of Economic Surveys*, 30(2):302–326, 2016.

Clark, T. E. and McCracken, M. W. Tests of equal forecast accuracy and encompassing for nested models. *Journal of Econometrics*, 105(1):85–110, 2001.

Clark, T. E. and West, K. D. Approximately normal tests for equal predictive accuracy in nested models. *Journal of Econometrics*, 138(1):291–311, 2007.

Commission. The pro-cyclicality of fiscal policy in EMU. Quarterly report on the euro area, European Commission, 2004.

Corsetti, G., Dedola, L., Jarociński, M., Maćkowiak, B., and Schmidt, S. Macroeconomic stabilization, monetary-fiscal interactions, and Europe's monetary union. *European Journal of Political Economy*, 57(C):22–33, 2019.

Dabla-Norris, E., Dallari, P., and Poghosyan, T. Fiscal spillovers in the euro area: Letting the data speak. Working Paper 17/241, International Monetary Fund, 2017.

Dao, M., Dizioli, A., Jackson, C., Gourinchas, P.-O., and Leigh, D. Unconventional fiscal policy in times of high inflation. IMF Working Papers 2023/178, International Monetary Fund, September 2023.

Darvas, Z. and Leandro, A. The limitations of policy coordination in the euro area under the European Semester. Bruegel Policy Contribution 2015/19, Bruegel, 2015.

D'Haultfoeuille, X. and Givord, P. La régression quantile en pratique. *Economie et Statistique*, 471(1):85–111, 2014.

Dolado, J. J., Maria-Dolores, R., and Naveira, M. Are monetary-policy reaction functions asymmetric? The role of non-linearity in the Phillips curve. *European Economic Review*, 49(2):485–503, 2005.

ECB. The impact of global value chains on the euro area economy. Occasional Paper Series 221, European Central Bank, April 2019.

ECB. Monetary-fiscal policy interactions in the euro area. Occasional Paper Series 273, European Central Bank, 2021.

Eichengreen, B. and Wyplosz, C. The Stability Pact: More than a minor nuisance? *Economic Policy*, 13(26), 1998.

Fatas, A. and Mihov, I. *Europe and the Euro*, chapter The Euro and Fiscal Policy, pages 287–324. The University of Chicago Press, 2010.

Forbes, K. J. Has globalization changed the inflation process? BIS Working Papers 791, Bank for International Settlements, 2019.

Forni, L. and Momigliano, S. Cyclical sensitivity of fiscal policies based on real-time data. *Applied Economics Quarterly*, 50(3):299–326, 2005.

Friedman, M. *Essays in positive economics.* University of Chicago Press, 1953.

Gagliardone, L., Gertler, M., Lenzu, S., and Tielens, J. Anatomy of the Phillips curve: Micro evidence and macro implications. NBER Working Papers 31382, National Bureau of Economic Research, Inc., June 2023.

Gali, J. and Gertler, M. Inflation dynamics: A structural econometric analysis. *Journal of Monetary Economics*, 44(2):195–222, October 1999.

Gavin, M. and Perotti, R. Fiscal policy in Latin America. NBER Macroeconomics Annual 1997, National Bureau of Economic Research, Inc., 1997.

Giacomini, R. and Rossi, B. Forecast comparisons in unstable environments. *Journal of Applied Econometrics*, 25(4):595–620, 2010.

Golinelli, R. and Momigliano, S. Real-time determinants of fiscal policies in the euro area. *Journal of Policy Modeling*, 28(9):943–964, 2006.

Golinelli, R. and Momigliano, S. The cyclical reaction of fiscal policies in the euro area: The role of modelling choices and data vintages. *Fiscal Studies*, 30(1):39–72, 2009.

Gootjes, B. and de Haan, J. Procyclicality of fiscal policy in European Union countries. *Journal of International Money and Finance*, 120(102276), 2022.

Gootjes, B. and de Haan, J. What makes discretionary counter-cyclical fiscal policy so difficult? An analysis of 32 OECD countries. *CESifo Economic Studies*, 69(1):1–20, 2023.

Gordon, R. J. U.S. Inflation, Labor's Share, and the Natural Rate of Unemployment. NBER Working Papers 2585, National Bureau of Economic Research, Inc., 1988.

Gordon, R. J. The Phillips curve is alive and well: Inflation and the NAIRU during the slow recovery. NBER Working Papers 19390, National Bureau of Economic Research, Inc., 2013.

Goujard, A. Cross-country spillovers from fiscal consolidations. *Fiscal Studies*, 38(2): 219–267, 2017.

Guerrieri, L., Gust, C. J., and Lopez-Salido, D. International competition and inflation: A New Keynesian perspective. *American Economic Journal: Macroeconomics*, 2(4): 247–280, 2010.

Harding, M., Lindé, J., and Trabandt, M. Understanding post-COVID inflation dynamics. *Journal of Monetary Economics*, 2023.

Haroutunian, S., Osterloh, S., and Slawińska, K. The initial fiscal policy responses of euro area countries to the COVID-19 crisis. Economic Bulletin Articles 1, European Central Bank, 2021.

Hazell, J., Herreño, J., Nakamura, E., and Steinsson, J. The slope of the Phillips curve: Evidence from U.S. states. *The Quarterly Journal of Economics*, 137(3):1299–1344, 2022.

Hobijn, B., Miles, R., Royal, J., and Zhang, J. The recent steepening of Phillips curves. Chicago Fed Letter 475, Federal Reserve Bank of Chicago (U.S.), 2023.

Holm-Hadulla, F., Hauptmeier, S., and Rother, P. The impact of numerical expenditure rules on budgetary discipline over the cycle. Working Paper Series 1169, European Central Bank, 2010.

Hooper, P., Mishkin, F. S., and Sufi, A. Prospects for inflation in a high pressure economy: Is the Phillips curve dead or is it just hibernating? *Research in Economics*, 74(1):26–62, 2020.

Hubrich, K. and Karlsson, T. Trade consistency in the context of the Eurosystem projection exercises. Occasional Paper Series 108, European Central Bank, 2010.

Ihrig, J., Kamin, S. B., Lindner, D., and Marquez, J. Some simple tests of the globalization and inflation hypothesis. *International Finance*, 13(3):343–375, 2010.

IMF. Cross-border impacts of fiscal policy: Still relevant? World Economic Outlook October, International Monetary Fund, 2017.

in't Veld, J. Fiscal consolidations and spillovers in the euro area periphery and core. Economic Papers 506, European Commission, 2013.

in't Veld, J. A public investment stimulus in surplus countries and its spillovers in the euro area. *National Institute Economic Review*, 239(1):53–62, 2017.

Judson, R. A. and Owen, A. L. Estimating dynamic panel data models: A guide for macroeconomists. *Economics Letters*, 65:9–15, 1999.

Kirsanova, T., Leith, C., and Lewis, S. W. Monetary and fiscal policy interaction: The current consensus assignment in the light of recent developments. *Economic Journal*, 119(541):482–496, November 2009.

Kloosterman, R., Bonam, D., and van der Veer, K. The effects of monetary policy across fiscal regimes. Working Papers 755, DNB, December 2022.

Koenker, R. *Quantile Regression.* Econometric Society Monographs. Cambridge University Press, 2005.

Koenker, R. and Bassett, G. Regression quantiles. *Econometrica*, 46(1):33–50, 1978.

Koenker, R. and Bassett, G. Robust tests for heteroscedasticity based on regression quantiles. *Econometrica*, 50(1):43–61, 1982.

Koenker, R. and Machado, J. A. F. Goodness of fit and related inference processes for quantile regression. *Journal of the American Statistical Association*, 94(448):1296–1310, 1999.

Koenker, R. and Xiao, Z. Inference on the quantile regression process. *Econometrica*, 70 (4):1583–1612, 2002.

Koester, G., Lis, E., Nickel, C., Osbat, C., and Smets, F. Understanding low inflation in the euro area from 2013 to 2019: Cyclical and structural drivers. Occasional Paper Series 280, European Central Bank, 2021.

Korobilis, D. Quantile regression forecasts of inflation under model uncertainty. *International Journal of Forecasting*, 33(1):11–20, 2017.

Lane, P. R. Inflation diagnostics. The ecb blog, European Central Bank, 2022.

Larch, M., Kumps, D., and Cugnasca, A. Fiscal stabilisation in real time: An exercice in risk management. *Economic Modelling*, 99, 2021a.

Larch, M., Orseau, E., and van der Wielen, W. Do EU fiscal rules support or hinder counter-cyclical fiscal policy? *Journal of International Money and Finance*, 112(C), 2021b.

Leeper, E. M., Walker, T. B., and Yang, S.-C. S. Fiscal foresight and information flows. *Econometrica*, 81(3):1115–1145, 2013.

Lopez-Salido, D. and Loria, F. Inflation at risk. Finance and Economics Discussion Series 2020-013, Board of Governors of the Federal Reserve System (U.S.), 2020.

Manzan, S. and Zerom, D. Are macroeconomic variables useful for forecasting the distribution of U.S. inflation? *International Journal of Forecasting*, 29(3):469–478, 2013.

Marinheiros, C. The stability and growth pact, fiscal policy institutions and stabilization in europe. *International Economics and Economic Policy*, 5(1):189–207, 2008.

Medel, C., Pedersen, M., and Pincheira, P. The elusive predictive ability of global inflation. Working Papers Central Bank of Chile 725, Central Bank of Chile, 2014.

Mikolajun, I. and Lodge, D. Advanced economy inflation: The role of global factors. Working Paper Series 1948, European Central Bank, 2016.

Mishkin, F. S. Globalization, macroeconomic performance and monetary policy. *Journal of Money, Credit and Banking*, 41(s1):187–196, 02 2009.

Moretti, L., Onorante, L., and Saber, S. Z. Phillips curves in the euro area. Working Paper Series 2295, European Central Bank, 2019.

Mourre, G., Astarita, C., and Princen, S. Adjusting the budget balance for the business cycle : The EU methodology. European Economy 536, European Commission, 2004.

Mundell, R. A theory of optimum currency areas. *American Economic Review*, 51: 657–665, 1961.

Musgrave, R. A. *The theory of public finance: a study in public economy.* McGraw Hill, New York, 1959.

Musso, A., Stracca, L., and van Dijk, D. Instability and nonlinearity in the euro area Phillips curve. Working Paper Series 811, European Central Bank, 2007.

Nickel, C., Koester, G., and Lis, E. Inflation developments in the euro area since the onset of the pandemic. *Intereconomics*, 57(2):69–75, 2022.

Nickell, S. Biases in dynamic models with fixed effects. *Econometrica*, 49(6):1417–1426, 1981.

Obstfeld, M. Harry Johnson's "Case for Flexible Exchange Rates" – 50 Years Later. NBER Working Papers 26874, National Bureau of Economic Research, Inc., March 2020.

Orphanides, A. Monetary policy rules based on real-time data. *American Economic Review*, 91(4):964–985, 2001.

Pesaran, H. M. General diagnostic tests for cross section dependence in panels. Cambridge Working Papers in Economics 0435, Faculty of Economics, University of Cambridge, 2004.

Pina, A. Elusive counter-cyclicality and deliberate opportunism? Fiscal policy from plans to final outcomes. Working Papers 6, Banco de Portugal, 2009.

Poghosyan, T. Cross-country spillovers of fiscal consolidations in the euro area. Working Paper 17/140, International Monetary Fund, 2017.

Poghosyan, T. and Tosun, M. S. Assessing activist fiscal policy in advanced and emerging market economies using real-time data. *Oxford Economic Papers*, 71(1):225–249, 2019.

Ramey, V. A. and Zubairy, S. Government spending multipliers in good times and in bad: Evidence from U.S. historical data. *Journal of Political Economy*, 126(2):850–901, 2018.

Reichlin, L., Ricco, G., and Tarbé, M. Monetary–fiscal crosswinds in the European Monetary Union. *European Economic Review*, 151(C), 2023.

Sargant, T. J. and Wallace, N. Some unpleasant monetarist arithmetic. Quarterly Review Fall, Federal Reserve Bank of Minneapolis, 1981.

Schmidt, K. The impact of monetary policy on the fiscal reaction : An empirical investigation in real-time. Unpublished manuscript, 2023.

Stock, J. H. and Watson, M. W. Phillips curve inflation forecasts. *Conference Series ; [Proceedings]*, 53, 2008.

Stock, J. H. and Watson, M. W. Modeling inflation after the crisis. NBER Working Paper 16488, National Bureau of Economic Research, Inc., 2010.

Tillmann, P. and Wolters, M. H. The changing dynamics of US inflation persistence: A quantile regression approach. *Studies in Nonlinear Dynamics and Econometrics*, 19(2): 161–182, 2015.

Woodford, M. Fiscal requirements for price stability. *Journal of Money, Credit and Banking*, 33(3):669–728, 2001.