Topics in Monetary Economics

by Selien De Schryder
### Doctoral Jury

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To my parents and Niels
Acknowledgements

There is no elevator to success. You have to take the stairs.
Anonymous

Completing a PhD does not just happen overnight. It involves a great deal of preparation, study, discipline and perseverance. Taking shortcuts is rarely an option and addressing criticism becomes an inevitable key aspect of your job. Fortunately, the life of a PhD candidate also has many attractive facets. Being able to build working skills and experience as a youngster in a dynamic and international environment is one of them. In fact, sharing skills and experiences is an unbearable aspect throughout a PhD. I am grateful to so many people for acquiring a combination of skills which were necessary to complete my PhD project. Here is the right place to thank them.

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I also appreciate the time and effort that the other members of my reading commission, Prof. dr. Gerdie Everaert (Ghent University), Prof. dr. Simon Price (Bank of England and City University London) and dr. Markus Eberhardt (University of Nottingham and Oxford University) have put into reading and commenting on my writings. Their detailed and thoughtful remarks and recommendations improved my papers to a considerable extent. My thanks also go out to prof. dr. Freddy Heylen, without whom I never would have considered to start a PhD in the first place and who was always ready for stimulating and constructive remarks and suggestions.

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Life of course does not entirely takes place at the office. I have the good fortune to be surrounded by friends and family who encouraged me throughout the PhD. I want to express my utmost gratitude to my parents who have supported me throughout every step in life. I would further especially like to thank my parents-in-law, my running coach Peter and my (former) team mates, my former class mates and my neighbors for their support, their understanding, the moments of fun and happiness, for just being there.

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Selien De Schryder,
September 2014
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Overview of the dissertation

As indicated by the title, the four papers in this dissertation tackle different topics in the field of monetary economics. The first chapter relates to global oil market dynamics, the second chapter highlights cross-country exports in the aftermath of the global financial crisis, the third chapter focuses on inflation dynamics in connection with bank credit evolutions whereas the final chapter concerns monetary regime dependence of the extent of wage indexation. Despite the wide angle of this dissertation, the chapters are very coherent because of the applied methodology. Each of the four chapters is characterized by the use of macro panels and accordingly by the application of macro panel data econometrics. Macro panels are particularly attractive as they allow to exploit both the cross section and time dimension of country-specific series but also require specific econometric care. Throughout the chapters in this dissertation, I have endeavored to exploit the advantages of the data for the topics in question while considering their peculiarities.

In the first chapter, which is joint work with Gert Peersman, we employ cross-country panel data to examine the role of the US dollar exchange rate as an economic driver of global oil demand. There is a growing consensus that global crude oil price fluctuations are mainly driven by changes in the demand for oil. Notwithstanding this consensus, the US dollar exchange rate has so far been ignored as an independent driver of oil demand in the empirical literature on global oil market dynamics. A similar argument holds for several empirical studies that exclusively focus on the analysis of the determinants of oil demand. This is surprising since global oil prices are predominantly expressed in US dollars. In this work, we explicitly recognize and quantify the role of the US dollar exchange rate as an economic driver of oil consumption and the possible consequences of its effect on the dynamics in the global crude oil market. This is a first contribution of the paper.

A second contribution of the paper to the literature is methodological. In particular, we (i) take the cointegration relationships between the variables into account by estimating
a panel error correction oil demand model, (ii) allow for cross-country heterogeneity of the coefficients which is present in the data, and (iii) correct for cross-sectional dependence in the error terms. We find that an appreciation of the US dollar real effective exchange rate leads to a decline in oil consumption in non-US dollar regions. Strikingly, the short-run US dollar exchange rate elasticity of oil demand turns out to be substantially larger than the elasticity of oil demand with respect to fluctuations in the global price of crude oil expressed in US dollar, i.e. more than double. A more detailed analysis of the pass-through of changes in global crude oil prices and the US dollar exchange rate to oil products end-user prices suggests that the difference in the magnitudes of both elasticities is the consequence of a significant larger pass-through of exchange rate fluctuations. A back-of-the-envelope calculation furthermore suggests that the US dollar exchange rate is an economically important contributor to the volatility of the global price of crude oil expressed in US dollar, due to its influence on oil demand. These findings underline that the US dollar exchange rate should be taken into account in the analysis of global oil market dynamics and sources of oil price fluctuations.

The second chapter of this dissertation is joint work with John Lewis and examines advanced economies’ export performance since the "Great Trade Collapse" (GTC) in a cross-country panel framework. Sharp falls in output in the immediate aftermath of the financial crisis were accompanied by even stronger falls in international trade. This observed plunge of global trade during late 2008 and early 2009 received the term "Great Trade Collapse" given its sudden nature and the extremely high level of synchronization across all advanced economies and nearly all industries. Our focus in this paper is to analyze what has happened to exports of advanced economies since the GTC and to gauge how consistent they were with predictions based on a pre-crisis panel model, rather than to uncover the cause of the GTC per se. In particular, we construct a forecast benchmark for the actual exports of a panel of advanced economies to analyse their export evolutions since the onset of the GTC. The capability of a pre-crisis standard export relationship to account for the evolutions of exports during and following the GTC is of great interest given its important policy implications. If the GTC marks structural changes in the relationship between trade performance and the traditional macroeconomic determinants, counterfactual export analyses based on traditional export determinants can provide misleading policy guidance.

We therefore construct a panel error correction model of goods exports for sixteen ad-
vanced economies and control for the possible existence of unobserved common correlated effects. We develop a novel measure of sectoral shifts in world trade and examine its effects next to the traditional price and income determinants of exports. We find the variable to have only a limited effect on export dynamics. We further find that the source of the real exchange rate shock matters for exports. The short-run response to relative unit labor costs is around six times as large as to nominal exchange rates.

We then use the model estimated over a pre-crisis sample period to undertake a forecast exercise to explore the dynamics of trade since the GTC. This exercise allows to construct a benchmark to which the advanced economies’ exports since the GTC can be compared. First, we assess how average post crisis export flows compare to the predictions of the pre-crisis model. We find that the pre-crisis model fares well in predicting the evolutions of exports over the entire forecast period once one controls for unobserved common factors. Second, we examine the forecasts for each country conditional on the country-specific variables and the common unobserved factors, proxied by the cross-section averages of all variables, to evaluate each country’s export performance against its peers. We find substantial variation across countries in terms of their actual exports relative to the forecasts based on the average panel coefficients. For the United Kingdom in particular, exports in 2011Q4 came in about 8 per cent below the benchmark suggested by international comparisons.

In the third chapter, inflation dynamics are analyzed in a cross-country Phillips curve framework while considering credit evolutions and the occurrence of financial stress to quantify the reaction of inflation to economic activity during different economic cycles. The immediate aftermath of the global financial crisis has been marked to contain a "missing disinflation puzzle". Historically, persistent and pronounced economic downturns gave rise to notable falls in the level of inflation but these falls are not observed during the severe economic downturns following the global financial crisis of 2008-2009. Nonlinearities and asymmetries in the Phillips curve relationship could offer an explanation for the muted reaction of inflation. The association of the downturns with a financial crisis however prompts one to question whether financial distortions have influenced the subsequent inflation dynamics beyond their impact via real output. The cross-country perspective in this work offers a way to obtain a sufficient number of observations on large negative output gaps, financial crises, credit downturns and their combinations.

We first analyze whether the reaction of inflation to economic activity differs depending
on the sign, magnitude and persistence of the deviation of output from its potential. This
analysis is warranted given the substantial evidence of nonlinearities and asymmetries in
the Phillips curve depending on the level of economic activity documented in the existing
literature and their potential to account for the missing disinflation puzzle. Second, we
examine whether the reaction of inflation to economic activity is in addition affected by
the credit cycle. We particularly concentrate on credit cycle downturns and upturns next
to the occurrence of banking crises. As such, we try to answer the question whether the
large extent of economic slack can explain the relatively mild disinflation during and in
the aftermath of the global financial crisis and whether the association with a financial
crisis and related credit evolutions attenuated the inflation reaction. In addition, we shed
light on the possibility of a speed limit effect during periods of spare capacity driven by a
bounce-back in output. The focus on the existence of a speed limit effect of spare capacity
on inflation is especially relevant nowadays as a speed limit would result in inflationary
pressures once the economy starts to recover, which is the situation currently faced by
most advanced economies.

Based on our analysis, we conclude that the mild inflation reaction subsequent to
the global financial crisis can be linked to an asymmetric reaction towards the extent of
spare production capacity whereas we do not find evidence that bank credit evolutions
and financial distress significantly alter the reaction of inflation to economic activity. We
further cannot find evidence that underpins the existence of a speed limit effect on inflation
when the extent of spare capacity shrinks. The fact that a long-lasting and pronounced
contraction is more likely to alter production resources obsolete and inadequate does not
seem to generate additional inflationary pressures.

The fourth chapter, which is joint work with Gert Peersman and Joris Wauters,
concerns the examination of the standard assumption in New Keynesian dynamic stoch-
astic general equilibrium (DSGE) models that wage indexation to past price inflation
is invariant to policy regimes. In particular, we examine the possibility of a link between
the degree of wage indexation and monetary policy through inflation uncertainty. Cross-
country panel data are employed to estimate a reduced-form empirical New Keynesian
wage Phillips curve where the degree of wage indexation to past inflation is allowed to
vary according to the monetary policy regime. Since an individual country’s monetary
policy regime is in general quite stable over time, a panel dataset approach allows us to
enhance the power of the test whether the degree of wage indexation depends on the type
of monetary policy regime as the number of observations increases significantly. We identify the monetary policy regime of a country in a specific period based on the presence of an explicit quantitative monetary target. Quantitative targets are transparent policy indicators and can be easily measured. A formal commitment to a quantitative target is therefore expected to improve the formation of inflation expectations and to reduce the inflation uncertainty of workers. The monetary target can take three forms: inflation, money growth and exchange rate targets. We distinguish between the presence of these three types of target, because the underlying dynamics of the strategies and the formation of inflation expectations are inherently different. Inflation targeting central banks for instance typically try to stabilize inflation in the short to medium term, whereas money growth targeting is more a commitment to low inflation in the long run. We further control for labor market institutions.

We find that wage indexation to past inflation varies across monetary policy regimes. Specifically, policy regimes that have an explicit quantitative inflation target are characterized by a lower degree of wage indexation. Discerning between three different types of explicit targets makes it clear that the extent of wage indexation is only significantly different (lower) in countries that have an inflation target whereas the effects of money and exchange rate targets are not significantly different from a regime without any formal quantitative target. These differences could be due to varying strengths of the nominal anchor under the different frameworks, as inflation targeting has been found to establish better anchored inflation expectations which, in turn, could strengthen the nominal anchor. Overall, our results question the structural nature of hard-wiring a fixed degree of wage indexation in standard DSGE models. Our work corroborates and extends the finding of substantial time variation in the degree of wage indexation for the US and it shows that the documented dependence of price indexation to monetary policy can be extended to wage indexation. From a policy standpoint, our findings suggest that counterfactual policy simulations and the analysis of optimal monetary policy based on modern macroeconomic models are potentially misleading.
Zoals aangegeven door de titel, behandelen de vier hoofdstukken van dit proefschrift verschillende domeinen van de monetaire economie. Het eerste hoofdstuk heeft betrekking tot de dynamiek van de globale oliemarkt, het tweede hoofdstuk legt de nadruk op de exporten van landen in de nasleep van de globale financiële crisis, het derde hoofdstuk focust op inflatie in combinatie met evoluties in de algemene kredietverlening terwijl het laatste hoofdstuk de afhankelijkheid van de mate van loonindexatie ten opzichte van monetaire regimes behandelt. Ondanks de ruime invalshoek van dit proefschrift, zijn de hoofdstukken sterk samenhangend omwille van de gehanteerde methodologie. Elk van de vier hoofdstukken wordt gekenmerkt door het gebruik van macro panels en bijgevolg door de toepassing van macro panel data econometrie. Macro panels zijn bijzonder attractief omdat ze toelaat om gebruik te maken van zowel de cross sectie als de tijdsdimensie van de landspecifieke reeksen maar ze vereisen ook een specifieke econometrische behandeling. Doorheen de hoofdstukken van dit proefschrift, heb ik gepoogd om deze voordelen van de data maximaal te gebruiken voor de behandelde thema’s terwijl ik de karakteristieken van de data in acht nam.

In het eerste hoofdstuk, dat een gezamenlijk werk is met Gert Peersman, hanteren we macro panel data om de rol van de wisselkoers van de Amerikaanse dollar in de globale olievraag na te gaan. Er is een toenemende consensus dat wijzigingen in de olieprijs voornamelijk gedreven worden door veranderingen in de vraag naar olie. Ondanks deze consensus is de wisselkoers van de Amerikaanse dollar tot nu toe genegeerd als onafhankelijke determinant van de vraag naar olie in de empirische literatuur over de wijzigingen in de globale oliemarkt. Dit is tevens aan de orde voor verscheidene empirische studies die zich enkel focussen op het analyseren van de determinanten van de vraag naar olieproducten. Dit is opmerkelijk aangezien de globale olieprijzen hoofdzakelijk zijn uitgedrukt in Amerikaanse dollar. In dit werk erkennen en kwantificeren we expliciet de rol van de
wisselkoers als een economische determinant van de consumptie van olie en de mogelijke
gevolgen van de effecten van de dollar op de globale oliemarkt. Dit is de eerste bijdrage
van de paper.

Een tweede bijdrage van de paper tot de betreffende literatuur is methodologisch van
aard. We nemen immers de coïntegrerende relatie van de variabelen in rekening door
het schatten van een "panel error correction" model. Daarnaast houden we rekening
met heterogene coëfficiënten over de landen heen en tot slot nemen we mogelijke cross-
sectie correlatie in de storingstermen in acht. Aldus vinden we dat een appreciatie van
de wisselkoers van de Amerikaanse dollar tot een daling leidt in de consumptie van olie
in niet-dollar landen. Opvallend daarbij is dat de korte termijn elasticiteit van de vraag
naar olie ten opzichte van de wisselkoers aanzienlijk groter blijkt te zijn dan de elasticiteit
ten opzichte van wijzigingen in de wereldprijs van ruwe olie en meer bepaald dubbel zo
groot. Een meer diepgaande analyse van het effect van wijzigingen in de wereldolieprijs en
wisselkoerswijzigingen op de prijzen van olieproducten voor de eindverbruiker geeft aan
dat de grootte van beide elasticiteiten gedreven wordt door een significant groter effect
van wisselkoerswijzigingen op de prijzen voor de eindverbruiker. De uitwerking van een
sterk vereenvoudigd vraag- en aanbod model leert ons tevens dat de wisselkoers van de
Amerikaanse dollar een economisch belangrijke bijdrage levert tot de volatiliteit van de
wereldolieprijs uitgedrukt in Amerikaanse dollar door de invloed op de vraag naar olie.
Deze resultaten benadrukken dat de Amerikaanse wisselkoers in acht dient genomen te
worden in de analyse van globale oliemarkt en wijzigingen in de olieprijs.

Het tweede hoofdstuk van dit proefschrift resulteert uit een samenwerking met John
Lewis en onderzoekt de prestaties van ontwikkelde landen inzake exporten sinds de "Great
Trade Collapse" (GTC) aan de hand van macro panel data. Zeer sterke dalingen van de
economische activiteit in de onmiddellijke nasleep van de financiële crisis gingen gepaard
met nog sterkere dalingen in de internationale handel. De geobserveerde daling in de
wereldhandel tijdens eind 2008 en begin 2009 kreeg de term "Great Trade Collapse"
onwille van de abruptheid en de uiterst hoge mate van synchronisatie over alle ontwikkelde
landen en quasi alle sectoren heen. De focus van deze paper ligt op de analyse van de
evoluties in de exporten van ontwikkelde landen sinds de GTC en op de samenhang van
deze evoluties met de voorspellingen op basis van modellen voorafgaand aan de GTC en
niet zozeer op het ontrafelen van de oorzaken van de GTC op zich. We construeren in
dit verband een maatstaf om de feitelijke exporten van een groep ontwikkelde landen te
beoordelen op basis van de voorspelling van ons model voor de aanvang van de GTC. Het vermogen van een standaard export vergelijking voorafgaand aan de crisis om de exporten tijdens en na de GTC te verklaren is van groot belang omwille van de beleidsimplicaties. Indien de GTC een structurele breuk kenmerkt in de relatie tussen exporten en de traditionele macro-economische determinanten, dan zullen de analyses op basis van de traditionele determinanten misleidende informatie geven.

We hanteren daarom een "panel error correction" model voor de exporten van goederen voor zestien ontwikkelde landen en controleren daarbij voor de mogelijke invloed van niet-geobserveerde gemeenschappelijke factoren. We stellen een nieuwe maatstaf samen voor sectorspecifieke wijzigingen in de wereldhandel en onderzoeken de effecten daarvan naast de traditionele determinanten van exporten, met name prijs en inkomen. Deze variabele blijkt slechts een beperkte invloed uit te oefenen op de dynamiek van exporten. Daarnaast vinden we dat de oorzaak van reële wisselkoerswijzigingen belangrijk is voor exporten. Op korte termijn is de reactie van exporten op een wijziging in relatieve arbeidskosten ongeveer zes keer zo groot als wijzigingen in de nominale wisselkoers.

We gebruiken vervolgens het model dat geschat is op de periode voorafgaand aan de crisis om een voorspelling te maken. Dit laat ons toe om een maatstaf op te stellen op basis waarvan de exporten van de individuele landen sinds de GTC kunnen geëvalueerd worden. Eerst gaan we na hoe de gemiddelde exporten sinds de GTC zich verhouden ten opzichte van de voorspelling door middel van de geschatte coëfficiënten op basis van de data voorafgaand aan de crisis. We vinden dat dit model de werkelijke daling in de gemiddelde exporten tijdens en na de GTC goed kan verklaren. Ten tweede analyseren we de exporten van elk individueel land conditioneel op de landspecifieke variabelen en de gemeenschappelijke factoren (die benaderd worden door de gemiddelden van de variabelen over de landen) om de prestaties inzake exporten van de landen ten opzichte van elkaar te vergelijken. Daarbij vinden we een aanzienlijke variatie tussen landen wat de relatieve verhoudingen tussen de werkelijke en voorspelde exporten op basis van de gemiddelde coëfficiënten betreft. Voor het Verenigd Koninkrijk in het bijzonder, waren de exporten ongeveer 8 procent lager dan de maatstaf op basis van de panel coëfficiënten.

Het derde hoofdstuk betreft een analyse van de dynamiek van inflatie aan de hand van een "Phillips curve" geschat over een panel van landen terwijl evoluties in de kredietverlening en het voorkomen van financieel moeilijke tijden in rekening worden genomen.
om de reactie van inflatie op de economische activiteit doorheen verschillende economische cyclussen na te gaan. De periode volgend op de globale financiële crisis werd gekenmerkt door een "missing disinflation" raadsel. Sterke en persistente economische recessies gaven in het verleden aanleiding tot aanzienlijke dalingen in het niveau van inflatie, maar dergelijke sterke dalingen zijn niet geobserveerd tijdens de sterke economische recessies volgend op de globale crisis van 2008-2009. Niet-lineariteit en asymmetrie in de "Phillips curve" relatie bieden een mogelijke verklaring voor de milde reactie van inflatie. De samenhang van de recessies met een crisis van financiële aard roept echter de vraag op of financiële verstoringen de daaropvolgende inflatie dynamiek kan beïnvloeden bovenop hun effect via het beïnvloeden van de reële output. Een panel data perspectief is in dit werk voornamelijk relevant aangezien het een mogelijkheid biedt om een voldoende aantal observaties van grote negatieve afwijkingen van output ten opzichte van het potentiële niveau van output, van financiële crisisissen, van dalingen in de kredietverstrekking en van hun combinaties te bekomen.

We onderzoeken eerst of de reactie van inflatie op de economische activiteit afhankt van het teken, de grootte en de persistentie van de afwijking van output van het potentiële niveau. Deze analyse is gerechtvaardigd gegeven het aanzienlijke bewijs van niet-lineariteit en asymmetrie in de "Phillips curve" afhankelijk van het niveau van economische activiteit in de bestaande literatuur en door het potentieel om het "missing disinflation" raadsel deels te verklaren. Daarnaast gaan we na of de reactie van inflatie op de economische activiteit eveneens wordt beïnvloed door de kredietcyclus en het voorkomen van een financiële crisis. Op die manier trachten we de vraag te beantwoorden of de grote mate van economische malaise de relatief milde deflatie tijdens en volgend op de globale financiële crisis kan verklaren en of de samenhang met een financiële crisis en de daarbij horende evoluties in de algemene kredietverstrekking door banken de reactie van inflatie temperden. Bovendien laten we ons licht schijnen op de mogelijkheid van "speed limit" effecten tijdens periodes van overcapaciteit gedreven door een heropleving van de geaggregeerde output. De focus op dergelijke effecten op inflatie is tegenwoordig uiterst relevant aangezien een snelheidslimiet in een opwaartse druk op inflatie zou resulteren van zodra de economie begint te heropleven en dit is exact de situatie waarin vele ontwikkelde economieën zich op dit moment bevinden.

Uit onze analyse kunnen we besluiten dat de milde reactie van inflatie na de globale financiële crisis kan gelinkt worden aan een asymmetrische reactie ten opzichte van de
capaciteit van de economie terwijl we geen bewijs vinden voor een significante invloed op de reactie van inflatie op de economische activiteit door evoluties in de kredietverlening of door algemene financieel moeilijke tijden. Tevens vinden we geen bewijs dat het bestaan van een "speed limit" effect op inflatie wanneer de mate van overcapaciteit verkleint onderbouwt.

Het vierde hoofdstuk, dat een gezamenlijk werk is met Gert Peersman en Joris Wauters, onderzoekt de basisveronderstelling in de Nieuw-Keynesiaanse dynamische stochastische evenwichtsmodellen (DSGE modellen) dat de indexatie van lonen aan prijzinflatie uit het verleden onafhankelijk is van het beleidsregime. We onderzoeken in het bijzonder de mogelijke link tussen de mate van loonindexatie en monetair beleid ten gevolge van onzekerheid omtrent inflatie. Macro panel data zijn toegepast om een gereduceerde empirische Nieuw-Keynesiaanse "wage Phillips curve" vergelijking te schatten waarbij de mate van loonindexatie aan inflatie uit het verleden kan wijzigen met het monetair beleidsregime. Aangezien het monetair beleidsregime van één enkel land in het algemeen vrij stabiel is doorheen de tijd, is een panel data methode belangrijk om voldoende observaties te bekomen die het schatten van de rol van een monetair beleidsregime mogelijk maken. We bepalen het monetair beleidsregime in een land op een bepaald tijdstip aan de hand van de aanwezigheid van een expliciete kwantitatieve monetaire doelstelling. Kwantitatieve doelstellingen zijn transparante beleidsindicatoren en kunnen op een eenvoudige manier worden nagegaan. Men verwacht daarom dat een formele verbinding tot een kwantitatieve maatstaf de vorming van inflatieverwachtingen verbetert en de onzekerheid over inflatie van de economische agenten reduceert. De monetaire doelstelling kan drie vormen aannemen: doelstellingen met betrekking tot het niveau van inflatie, de geldgroei of de wisselkoers. We maken een onderscheid tussen deze drie types van doelstellingen aangezien de onderliggende strategieën en de manier van het vormen van verwachtingen intrinsiek verschillend zijn. Centrale banken met een inflatie doelstelling proberen bijvoorbeeld inflatie te stabiliseren op korte tot middellange termijn terwijl centrale banken met een doelstelling voor de geldgroei zich meer engageren tot een lage inflatie op lange termijn. We controleren tevens voor evoluties in arbeidsmarktinstellingen.

We vinden dat loonindexatie aan prijzinflatie uit het verleden varieert onder monetaire beleidsregimes. Regimes met een expliciete kwantitatieve inflatie doelstelling worden gekenmerkt door een lagere mate van dergelijke loonindexatie. Wanneer we een onderscheid maken tussen de drie verschillende types van expliciete doelstellingen, dan blijkt de
mate van loonindexatie enkel significant verschillend (lager) te zijn voor landen met een
doeleindeling voor inflatie terwijl de effecten van de doelstellingen met betrekking tot de
geldgroei en de wisselkoers niet significant verschillen van een regime zonder enige kwantitatieve doelstelling. Deze verschillen zijn mogelijk het gevolg van een verschil in de
sterkte van de monetaire verankering onder de verschillende monetaire structuren. Een
rechtstreekse doelstelling voor inflatie werd immers reeds gelinkt aan een betere verankering
van de inflatieverwachtingen, wat op zijn beurt leidt tot een betere verankering van het
monetaire doel. In het algemeen stellen onze resultaten de inherente structurele aard van
de mate van loonindexatie in de standaard DSGE modellen in vraag. Ons werk bevestigt
en breidt de reeds gedocumenteerde bevindingen van een substantiële mate van variatie in
de mate van loonindexatie over de tijd voor de Verenigde Staten uit en het toont aan dat
de eveneens reeds aangetoonde afhankelijkheid van prijsindexatie tot monetaire regimes
kan uitgebreid worden tot loonindexatie. Vanuit een beleidsoogpunt, suggereren onze re-
sultaten dat beleidssimulaties en analyses van een optimaal monetair beleid gebaseerd op
de moderne macro-economische modellen potentieel misleidend zijn.
CHAPTER 1

The U.S. Dollar Exchange Rate and the Demand for Oil

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Abstract

Using recent advances in panel data estimation techniques, we find that an appreciation of the US dollar exchange rate leads to a significant decline in oil demand for a sample of 65 oil-importing countries. The estimated effect turns out to be considerably larger than the impact of a shift in the global crude oil price expressed in US dollar. This finding appears to be the consequence of a stronger pass-through of changes in the US dollar exchange rate to domestic end-user oil products prices relative to changes in the global crude oil price. Furthermore, we demonstrate the relevance of US dollar fluctuations for global oil price dynamics.

JEL classification: C33, F31, Q41

Keywords: Oil demand, US dollar exchange rate, oil price pass-through, panel data

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1 Introduction

There is a growing consensus that global crude oil price fluctuations are mainly driven by changes in the demand for oil. Hamilton (2009), for instance, argues that strong growth in world income was the primary cause of the oil price surge in 2007-08, whereas the subsequent dramatic collapse of oil prices was the result of the global economic downturn in the aftermath of the financial crisis. Furthermore, Peersman (2005), Kilian (2009), Peersman and Van Robays (2009), Lombardi and Van Robays (2011) and Kilian and Murphy (2012) disentangle different sources of oil price shocks within a structural vector autoregressive (SVAR) model, and find a dominant role for shocks at the demand side of the global crude oil market. In order to better understand oil market fluctuations, a more detailed analysis of the drivers of oil demand is thus desirable.

In this paper, we examine the role of the US dollar exchange rate for oil consumption. The US dollar exchange rate has so far been ignored as an independent driver of oil demand in the empirical literature on global oil market dynamics. This is surprising since global oil prices are predominantly expressed in US dollars. According to the local oil price channel, a shift in the dollar exchange rate should then affect the demand for crude oil in countries that do not use the US dollar for local transactions (Austvik 1987). For instance, when the US dollar exchange rate depreciates, oil becomes less expensive in local currency for consumers in non-US dollar regions, boosting their demand for oil. The rise in oil demand for countries that do not use the dollar for local transactions should in turn influence global oil production and oil prices expressed in US dollar. This line of reasoning was raised in the work of Brown and Philips (1984) and Huntington (1986), and is supported by the data shown in Figure 1.1. The panels in the figure show the evolution of the real effective US dollar exchange rate, as well as the deviation of oil consumption from its trend, for a set of countries (and country aggregates) that are examined in this paper. As can be seen in the figure, an appreciation (depreciation) of the dollar exchange rate is often accompanied by a decline (rise) in oil consumption relative to its trend evolution, indicating a fall (rise) in oil demand. Shifts in the US dollar exchange rate could thus be important for global oil market dynamics.

A similar argument holds for several studies that exclusively focus on the analysis of the determinants of oil demand. In particular, Gately and Huntington (2002), Cooper (2003), Dargay, Gately and Huntington (2007), Narayan and Smyth (2007) and Dargay
and Gately (2010) amongst others estimate oil demand functions for multiple countries. These studies consider oil demand as a positive function of income per capita and a negative function of its own price. For the latter, they typically use global crude oil prices expressed in US dollars due to the lack of sufficient and/or reliable data on local oil prices. The influence of shifts in the US dollar exchange rate on oil demand is hence not taken into account. Some studies (e.g. Griffin and Schulman 2005; Dargay et al. 2007; Dargay and Gately 2010; Fawcett and Price 2012) do use local oil/gasoline prices in the estimations, but do not distinguish between local oil price movements caused by global oil price shifts and movements caused by changes in the value of the US dollar. There is, however, no a priori reason to assume that the pass-through and influence of both sources of oil price shifts on oil demand is the same.

We formally investigate the effects of shifts in the US dollar exchange rate on oil demand in non-US dollar regions, by estimating the determinants of oil consumption per capita for a panel of 65 oil-importing countries over the sample period 1971-2008. A panel data approach is commonly used in the literature on oil (energy) demand, as it allows to exploit both the cross section and the time dimension of the data. We conduct panel estimations for respectively a sample of 23 OECD countries, 42 non-OECD countries and all 65 oil-importing countries. Besides real GDP per capita, we include global real crude oil prices expressed in US dollar, as well as the real US dollar exchange rate in the estimations. An explicit analysis of the role of the US dollar as a possible driver of oil consumption is a first contribution of the paper.

A second contribution of the paper is methodological. In particular, most existing panel data studies on oil demand do not fully take into account the specific salient features of macro panel data sets such as heterogeneity of the coefficients, unit root behavior and cross-country dependence, even though the neglect of these matters can result in misleading estimation outcomes. We apply recent advances in panel estimation techniques that are

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2To our knowledge, the only empirical study which also considers the US dollar exchange rate as a possible determinant of oil demand is Askari and Krichene (2010). They estimate, however, a time series simultaneous equation model for (aggregate) world oil demand and supply between 1970 and 2008, whereas we estimate the impact of the US dollar exchange rate for a panel of 65 countries. In addition, they examine the effect of the exchange rate as part of a monetary policy channel affecting global oil prices, rather than an independent driver of oil demand.
capable to handle these econometric issues. Specifically, we (i) take into account the long-run relationship between the variables by estimating a panel error correction oil demand model, (ii) allow for cross-country heterogeneity of the coefficients which is present in the data, and (iii) consider cross-sectional dependence in the error terms. The application of these econometric advances and the addition of the US dollar exchange rate as a driver of oil consumption turn out to matter for some of the estimated elasticities.

We find that an appreciation of the US dollar real effective exchange rate leads to a decline in oil consumption in non-US dollar regions. Strikingly, the short-run US dollar exchange rate elasticity of oil demand turns out to be substantially larger than the elasticity of oil demand with respect to fluctuations in the global price of crude oil expressed in US dollar. A more detailed analysis of the pass-through of changes in global crude oil prices and the US dollar exchange rate to oil products end-user prices for a subset of 20 OECD-countries suggests that the difference in the magnitudes of both elasticities is the consequence of a significant larger pass-through of exchange rate fluctuations. A back-of-the-envelope calculation furthermore suggests that shifts in the US dollar exchange rate are an economically important contributor to the volatility of the global price of crude oil expressed in US dollar, due to its influence on oil demand. These findings underline that the US dollar exchange rate should be taken into account in the analysis of global oil market dynamics and sources of oil price fluctuations.

The remainder of this paper is organized as follows. In the next section, we describe the baseline empirical model for oil demand and discuss some econometric issues. Section 4 discusses the estimation and robustness of the results. The pass-through of changes in global oil prices and the real effective US dollar exchange rate to local end-user oil prices is examined in section 4, while the economic relevance of the US dollar exchange rate for global oil market dynamics is assessed in section 5. Finally, section 6 concludes.

## 2 Empirical oil demand model

In this section, we describe the benchmark oil demand model that will be used in the estimations. Our sample contains 65 oil-importing countries that do not have the US dollar as their local currency and covers the period 1971-2008. Details of the data and

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3 The United States is hence not included in the analysis, which should be taken into account when comparing the results with existing studies. We do also not consider oil-exporting countries, since oil de-
a list of the countries can be found in Appendix A. Consider the following general oil demand specification for country $i$ at time $t$:

$$dem_{it} = f(gdp_{it}, oilp_t, rer_t, trend_t, c_i)$$  \hspace{1cm} (1.1)

where $dem_{it}$ is total oil consumption per capita, $gdp_{it}$ real income per capita, $oilp_t$ the world real US dollar crude oil price and $rer_t$ the real effective US dollar exchange rate, $trend_t$ a linear time trend and $c_i$ a country-specific constant. All variables are converted to natural logarithms, such that the model is of the constant elasticity form. The data are at annual frequency.

The existing empirical literature typically considers oil consumption, or energy consumption more generally, as a positive function of real income and a negative function of its own price (e.g., Dahl and Sterner 1991; Dahl 1993; Espey 1998; Gately and Huntington 2002; Cooper 2003; Griffin and Schulman 2005; Hughes, Knittel and Sperling 2008; Lee and Lee 2010; Dargay and Gately 2010). In line with these studies, we include country-specific real GDP per capita in the general oil demand specification. Real GDP is assumed to represent the energy-using capital stock, such as buildings, equipment and vehicles (Dargay and Gately 2010).

As a measure for the own price of oil demand, most studies use the global price of crude oil expressed in US dollar (supra, page 3). However, as we have explained in the introduction, the price that consumers face in countries that do not use the US dollar as a currency for local transactions, is the price of oil determined in dollars multiplied by the country’s exchange rate against the US dollar, i.e. the number of units of national currency needed to buy one US dollar (Austvik 1987). Some studies use local oil/gasoline prices in the estimations (supra, page 3), but this does not allow to distinguish between local oil-price shifts caused by changes in the global price of crude oil, or changes in the US dollar exchange rate, which is the central research question in this paper. Moreover, the lack of availability of country-specific end-user prices would constrain the sample considerably.\footnote{Country-specific end-user prices are only available for a limited number of OECD countries, which limits the usefulness for the analysis of global oil market dynamics, in particular since non-OECD countries constitute an increasingly large share of this market. The use of global oil prices also avoids endogeneity problems of using local oil prices. Specifically, in contrast to local oil (gasoline) prices, it is more plausible to assume that global crude oil prices are exogenous for an individual country’s oil demand.}

mand in these countries has been found to behave very differently. See for example Gately and Huntington (2002).
Accordingly, we include the global real crude oil price expressed in US dollar, as well as the real US dollar effective exchange rate as two separate variables in our empirical oil demand model.

We use the US dollar real effective exchange rate rather than real bilateral exchange rates in the benchmark estimations for three reasons. First, bilateral exchange rates (or domestic CPI) are not available for several countries over the whole sample period, which would reduce the size of the dataset. Second, bilateral exchange rates suffer an endogeneity problem as the demand for oil of an individual country is expected to influence its own exchange rate. Given the moderate weight of each individual country in the US trade basket, this is much less the case for the US dollar effective exchange rate. Third, a multilateral weighted exchange rate is more useful to examine the role of changes in the US dollar for global oil market dynamics. In section 3.2, however, we assess the robustness of the results for a specification with the bilateral exchange rates estimated for a subsample of countries using instrumental variables.

Finally, microeconomic theory (e.g. Mas-Colell, Whinston and Green 2007) suggests that oil demand is also a function of the economy’s structure, technology and the prices of substitutes. To capture the former two, all our estimations contain a country-specific constant and a linear trend. In addition, we add proxies for unobserved common factors, as discussed in section 2.3. Unfortunately, prices of substitutes are not available for our sample period. Griffin and Schulman (2005), however, report only weak substitution effects when they include the real price of substitute fuels in a demand model for petroleum products, and argue that omitting cross-price effects does not appreciably alter their results.5

2.1 Panel unit root and cointegration tests

To avoid spurious regression results, we first examine the time series properties of the variables. Since cross section dependence tests (CD tests, Pesaran 2004) on the residuals of Augmented Dickey Fuller regressions (ADF regressions, Dickey and Fuller 1979) indicate

5Frankel (2006) argues that oil and other commodity price developments are influenced by interest rates. Specifically, when the interest rate declines, commodities become more attractive as an asset for investors. In addition, a lower interest rate stimulates overall demand, including the demand for oil. Notice, however, that this is not relevant for our analysis since we consider oil consumption (not inventories) at the LHS of the oil demand function, while real GDP is included at the RHS.
a significant degree of cross section dependence for the country-specific variables \((dem_{it} \text{ and } gdp_{it})\), we employ the Panel Analysis of Non-stationarity in Idiosyncratic and Common components (PANIC) test proposed by Bai and Ng (2004) for both series.\(^6\) The number of common factors \((r)\) is determined by the Bai and Ng (2002) criteria.\(^7\) Table 1.1 shows that \(r\) varies between one and four, depending on the variable and the criterion under consideration. Both series are found to be non-stationary for all specifications, which is due to the non-stationarity of the common component and the idiosyncratic component (or solely to the former).

Given that the global real crude oil price variable and the US dollar real effective exchange rate are observed common factors, we use standard ADF tests for both series. For global crude oil prices, the existence of a unit root cannot be rejected for both a constant only and linear trend model. This is, however, not the case for the US dollar real effective exchange rate, for which the ADF test rejects non-stationarity of the series. This finding is at odds with the empirical purchasing power (PPP) literature, where standard univariate ADF tests typically fail to reject the null hypothesis. Engel (2000) shows that standard unit root tests may, however, be biased in favour of rejecting non-stationarity if the real exchange rate has a stationary and a non-stationary component. For this reason, and to ensure consistency with the other variables in the model, we continue to treat the US dollar real effective exchange rate as a non-stationary variable in the analysis.

In a second step, we test for a long-run relationship amongst the variables using the panel error correction test of Gengenbach, Urbain and Westerlund (2008), henceforth GUW test. The test is based on the significance of the error correction term in the panel error correction model (ECM). Compared to residual-based panel unit root tests, the GUW test has the advantage that it is not subject to the common factor critique (Kremers, Ericsson and Dolado 1992) and that it does not rely on a stepwise testing procedure.\(^8\) Notice that the GUW test is nevertheless more restrictive than residual-based tests by imposing weak exogeneity on the country-specific regressors of the ECM.

\(^6\)Other panel unit root tests that also use a common factor representation of the data to allow for cross-section dependence (Moon and Perron 2004; Pesaran 2007) impose restrictions on the number of common factors and/or assume stationarity of the common factors. Given the results concerning the number and the stationarity properties of the common factors, these alternative tests are not used.

\(^7\)We consider the IC1, IC2 and BIC3 criteria. The BIC3 criterion is more robust when there is cross correlation in the idiosyncratic errors (Bai and Ng 2002).

\(^8\)The common factor critique applies to residual-based panel cointegration tests as they rely on residual rather than structural dynamics (Gengenbach et al. 2008).
and strong exogeneity on the common factors, whereas residual-based tests allow for full endogeneity. To take this restriction into account, we have also applied residual-based panel cointegration tests in the spirit of Banerjee and Carrion-i-Silvestre (2006) to check the robustness of the results. The results of the tests are shown in Table 1.2. The pooled GUW tests reject the null hypothesis of no error correction between \( \text{dem}_{it}, \text{gdp}_{it}, \text{oilp}_{t} \) and \( \text{rer}_{t} \) for the model under consideration, i.e. including a constant and trend. The alternative residual-based cointegration tests confirm this result. As a consequence, we can safely conclude that \( \text{dem}_{it}, \text{gdp}_{it}, \text{oilp}_{t} \) and \( \text{rer}_{t} \) are cointegrated at the panel level.

### 2.2 Panel error correction oil demand model

Having established cointegration between the variables, we can formulate our general oil demand specification as a panel ECM:

\[
\Delta \text{dem}_{it} = \alpha_i + \tau_i \times \text{trend}_t + \lambda_i \times \text{dem}_{i,t-1} + \gamma_i \times \Delta \text{gdp}_{i,t-1} + \beta_i \times \text{oilp}_{t-1} + \theta_i \times \Delta \text{rer}_{t-1} + \\
\gamma_i \times \Delta \text{gdp}_{it} + \beta_i \times \Delta \text{oilp}_{t} + \theta_i \times \Delta \text{rer}_{t} + \varepsilon_{it} \tag{1.2}
\]

Equation (1.2) is the baseline empirical specification for the panel ECMs that will be estimated in this paper. Gately and Huntington (2002) and Griffin and Schulman (2005) are most closely related to our study as they both estimate single equation total oil demand models for a panel of multiple countries with a moderate time dimension. Before we estimate the panel ECM, it is important to discuss two econometric issues which are generally disregarded in the existing oil demand literature, namely slope heterogeneity and cross-sectional error dependence.

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9 The approach we take to examine the robustness of the GUW test results is the following: we apply the continuously-updated and bias-corrected (CupBC) estimator of Bai, Kao and Ng (2009) to the long-run cointegration equation and we test the (de-factored) residuals for a unit root using the PANIC test procedure.

10 The lag order of the dynamic adjustment process is imposed to be 0 for all variables for reasons of parsimony. Experiments with more lags, however, do not alter the results.

11 Gately and Huntington (2002) have a sample of 93 countries over 1971-1997. Griffin and Schulman (2005) use data on 16 OECD countries over 1961-1999. Griffin and Schulman (2005) include retail oil prices instead of the world crude oil price in their model, which makes their estimation outcomes less adequate to serve as a benchmark for the baseline model. Both works consider the standard Fixed Effects (FE) estimator. Notice that this FE estimator suffers small \( T \) problems in dynamic panels (Arellano and Bond 1991). Given that the time dimension of our sample is moderately large, we assume that this bias is not relevant for our purposes. Nickell (1981) shows that the upward bias on the error correction term becomes insignificant when \( T \to \infty \).
2.3 Econometric issues

In order to obtain reliable estimates, we need to consider two important econometric features of macro panel data. The first one concerns heterogeneity in the slope coefficients. Heterogeneity in the slope coefficients renders the standard FE estimator biased as the latter assumes homogeneity in the slope coefficients. Some studies, e.g. Gately and Huntington (2002) and Dargay and Gately (2010), notice a substantial heterogeneity within non-OECD countries and split their sample in different groups of countries, i.e. OECD countries, oil exporters, income growers and other non-OECD countries. Given the cross-country differences in economic structures, the assumption of homogenous slope coefficients within groups is nevertheless questionable, including the group of OECD countries. Indeed, the application of Swamy’s Wald test consistently reveals that the homogeneity restriction on the slope coefficients is not valid, even for the subsample of OECD countries (see Table 1.3). The FE estimates are hence potentially misleading. Accordingly, we use the Mean Group (MG) estimator in the analysis, which offers a consistent alternative as the MG estimator does not impose homogeneity.\footnote{Since the MG estimator requires large $N$, and allows for cross-country heterogeneity anyway, we pool all non-OECD countries in one group. A further decomposition of the non-OECD countries in e.g. fast-growing and income-growth stagnating countries, as in Gately and Huntington (2002) and Dargay and Gately (2010), might be interesting, but is unfortunately not feasible due to the limited number of income-growers in the sample.}

The second important feature of macro panel data estimations relates to error cross section dependence. In particular, the results of standard FE and MG estimators are inconsistent and have biased standard errors when the observed explanatory variables are correlated with unobserved common factors (Pesaran 2006). For oil demand, this is likely the case. Country-specific income, the real price of crude oil, as well as the US dollar real effective exchange rate could for instance be driven by a common global business cycle. The existing empirical oil demand studies do not consider the potentially far-reaching consequences of cross-sectional dependence. The presence of cross section dependence in the error terms of dynamic models could be tested by means of the CD test of Pesaran (2004). Applying the CD test to our panel error correction oil demand model shows that there is a significant degree of cross-sectional correlation in the error terms for both the FE and MG estimators (see Table 1.3), which confirms the need to attempt to take into account the dependence.
We therefore apply the Bai and Ng (2004) PANIC decomposition to the residuals of the model in order to estimate the common components in the residuals. In the spirit of Bai et al. (2009), the estimated common factor(s) is (are) then in a second step included in the model to get consistent estimates. This procedure allows us to remove, or at least significantly reduce, the common factors that are present in the residuals of the first step.\textsuperscript{13} Another advantage is that possible non-linear unobserved common variables such as technological change (as in Griffin and Schulman 2005) can be appropriately controlled for without imposing an homogeneous coefficient. By including the estimated common components of the residuals of the MG regression as a proxy for omitted common variables in the model, we notice a substantial decline of the cross-sectional correlation in the residuals for the sample of OECD countries (Table 1.3, last two lines). The effect for the non-OECD group is, in contrast, limited.

In sum, in contrast to the existing empirical evidence on the demand for oil, we do not only examine the role of the US dollar exchange rate for the demand for oil, we also apply panel estimators that take both heterogeneity of the coefficients and cross-sectional dependence into account.

3 Empirical results

3.1 Panel estimations

Table 1.3 summarizes the estimation results of the panel error correction model as described in section 2.1. In order to compare with the existing evidence, we report the results for respectively OECD countries, non-OECD countries and the total sample of oil-importing countries. For each sample, we show the results for the FE, MG and MG estimator adjusted for cross-sectional dependence (MG\_Ft), which should allow us to evaluate the relevance of the econometric features discussed in section 2.3 for the estimation results. Notice that all estimated income and oil price coefficients reported in the paper are very similar when we re-estimate the oil demand models without the exchange rate

\textsuperscript{13}The drawback of this approach, however, is that the estimation error from the first step carries over to the subsequent steps. The presence of multiple observed common factors as explanatory variables in our model makes the more standard application of the Common Correlated Effects (CCE) estimators of Pesaran (2006) to eliminate cross-sectional dependence however unattractive.
variable. The corresponding conclusions are thus robust for the inclusion of the exchange rate (unless otherwise mentioned).

**Income elasticity**  We consistently find a significant positive effect of real GDP on the demand for oil. The short-run income elasticity in OECD countries is 0.69, which is larger than the 0.40 found by Griffin and Schulman (2005). Furthermore, the estimated average impact of economic activity on the demand for oil is similar for non-OECD countries (0.53) and the total sample of countries (0.60). The long-run income elasticity coefficients are, in contrast, more diverse across both groups of countries. Specifically, the average long-run income elasticity turns out to be 0.52, 1.06 and 0.94 for respectively OECD, non-OECD and all 65 oil-importing countries, which is in line with most existing studies. A lower income elasticity in more developed countries is, for instance, also found by Gately and Huntington (2002).

The econometric issues that we discussed in section 2.3 seem to matter for the magnitudes of the estimates. Specifically, the long-run income elasticity for OECD countries increases from 0.52 to 0.92 if we do not take into account cross-sectional dependence in the error terms, and even to 1.11 if we also do not allow for cross-country heterogeneity in the coefficients. Interestingly, exactly the opposite happens for non-OECD countries, i.e., the long-run income elasticity declines from 1.06 to respectively 1.02 and 0.73 when there is not allowed for correlation and heterogeneity across countries. In other words, the bias resulting from the use of a FE estimator can be relevant and could work in both directions.

**(Global) oil price elasticity**  There are numerous papers that estimate the effects of a shift in (global) crude oil prices on the demand for oil. Most studies report a relatively low, or even an insignificant (e.g. Askari and Krichene 2010) short-run price elasticity of oil demand, which is important because a low oil price elasticity implies that any disruption

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\(^{14}\)Notice that the baseline specifications of Gately and Huntington (2002) and Griffin and Schulman (2005) are different from our baseline model. In particular, Gately and Huntington (2002) allow for asymmetric responses to increases and decreases in the crude oil price, whereas Griffin and Schulman (2005) add a time effect to capture the non-linear nature of technological change and show that symmetric price responses cannot be rejected once one allows for a time effect. Adding proxies for unobserved common factors (as in MG\(_{Ft}\)) is equivalent to the approach of Griffin and Schulman (2005), with more degrees of freedom. The possibility of asymmetric price-responsiveness is further examined in the next section.
in oil production has a considerable impact on the price of oil. As can be seen in Table 1.3, we find a significant negative (short-run) effect of a change in the global price of crude oil expressed in US dollars on the demand for oil in OECD-countries (-0.05), non-OECD countries (-0.03) and the overall sample of countries that do not use the US dollar for local transactions (-0.04). These short-run elasticities turn out to be very similar for the different estimators and are in line with several other panel studies.\textsuperscript{15} We further find a stronger response of oil demand to a global oil price shift in the long run, although the magnitude of the long-run coefficients are much lower for the MG coefficients than for the FE estimates. The estimated long-run price elasticities are respectively -0.12 and -0.15 for OECD and non-OECD countries.

**Exchange rate elasticity** Our results reveal that there is a strong effect of the US dollar exchange rate on oil demand in the rest of the world, despite the fact that we control for country-specific real GDP and global crude oil prices, which supports the conjecture that the US dollar exchange rate is a significant driver of oil demand. More specifically, when the US dollar real effective exchange rate appreciates by 1 percent, there is a short-run decline in oil demand of 0.19 percent in OECD countries. Strikingly, the estimated elasticity is considerably bigger than the global crude oil price elasticity expressed in US dollar. The equality of the short-run price and real exchange rate elasticity is rejected at the panel level.\textsuperscript{16} The negative effect of the exchange rate on oil demand in OECD countries rises even further to -0.31 in the long run, which is also much larger than the long-run elasticity of the global oil price determined in US dollar of -0.12.

The short-run impact of the US dollar exchange rate on oil demand in non-OECD countries is much lower (-0.06) and statistically not significant. Notice that the latter is not the case for the FE estimator that is typically used in the oil demand literature, which confirms that not taking into account the features of macro panel data sets could be misleading for the interpretation of the results. A possible explanation for the insignificant exchange rate elasticity coefficient can be the characteristics of the group of non-OECD

\textsuperscript{15}Larger magnitudes for the short-run oil price coefficient are found by Bodenstein and Guerrieri (2011) within a DSGE framework, and in the SVAR studies of Baumeister and Peersman (2013) and Kilian and Murphy (2012).

\textsuperscript{16}Askari and Krichene (2010) estimate a time series simultaneous equation model for global oil demand and supply using quarterly data over a similar sample period (1970-2008) and also find an impact of the US dollar (nominal) exchange rate on oil demand which is stronger in magnitude than the effect of the price of oil, but both elasticities turn out to be insignificant.
countries. Specifically, some of the non-OECD countries had varying exchange rate regimes over time or experienced exchange rate crises during the sample period, which could reduce the estimated response of oil consumption to US dollar fluctuations. The estimated long-run exchange rate elasticity coefficient is, however, significant and about twice the size of the long-run price elasticity. This indicates that the factors that prevent oil consumption in these countries to respond to changes in the value of the US dollar diminish over time. Finally, the US dollar real effective exchange rate elasticity for the total sample of 65 oil-importing countries is -0.09 and significant, which is again more than double the global oil price elasticity expressed in US dollar for the same sample of countries. In sum, the US dollar exchange rate matters for oil demand in countries which do not use the dollar as a currency for local transactions. A weakening of the US dollar boosts oil consumption in these countries. In the next subsection, we assess the robustness of this novel finding, while we examine the relevance for global oil market dynamics in section 5.

3.2 Robustness checks

In this section, we assess the robustness of the baseline results. We first check whether the estimated exchange rate elasticities are robust to the choice of a price-symmetric model by allowing for asymmetric oil price reactions. We then examine the robustness of the results for possible endogeneity problems between oil demand and respectively global crude oil prices and the US dollar exchange rate.

Asymmetric-price model The possibility of asymmetric responses of oil consumption to price changes has received attention in several empirical studies. The underlying idea is that higher prices induce more investment in energy-efficient equipment and retrofitting of existing capital. When prices fall, however, there is no switch back to less-efficient capital, although there could be more intensive usage (Griffin and Schulman 2005). More recently, Dargay et al. (2007) and Dargay and Gately (2010) even allow for a different reaction of oil demand to price increases that result in a new historical maximum price \( p_{\text{max}} \), to price increases back to the previous maximum \( p_{\text{pre}} \), and to price decreases

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17 Some countries temporarily had a fixed and/or crawling peg exchange rate regime. The sample, however, does not contain countries which had a fixed peg to the US dollar over the entire sample period.

They find that oil demand responds differently to the different price shifts, with the largest (negative) effect of price increases that result in new maximum values.\footnote{These studies also allow for possible asymmetric responses to income changes, but Dargay and Gately (2010) point out that this approach is primarily appropriate for oil exporting countries, a group which is not included in our analysis. We therefore do not extend the model with this type of asymmetry.}

Table 1.4 shows the results when we extend our baseline specification with the price decomposition of Dargay et al. (2007) and Dargay and Gately (2010). Using Wald tests at the panel level, we consistently find that the effect of a price increase to a new maximum is significantly larger than both other price movements, whereas there are no significant differences between a price increase back to an earlier maximum and a price cut. More importantly, however, we find that the exchange rate elasticity of oil demand is still significant and similar in magnitude to the benchmark results reported in section 4 and Table 1.3. Interestingly, a Wald test cannot reject the hypothesis that the responsiveness of oil demand to changes in the US dollar exchange rate and a price increase to a new maximum are significantly different. Overall, we can conclude that the benchmark results are robust when we allow for asymmetric oil price responses of oil consumption. The US dollar exchange rate is still a significant driver of oil consumption, with an impact that is generally larger than shifts in global crude oil prices expressed in US dollar. In the rest of the paper, we therefore continue to use the price-symmetric specification, with the estimated unobserved common factors acting as a proxy for technological change.

**Instrumental variables estimation** In line with the existing cross-country panel studies on oil demand, we have assumed in the benchmark estimations that the demand for oil of an individual country does not influence the global price of crude oil on impact. However, if a country has a large share in global oil consumption, the assumption of exogenous oil price movements could be violated.\footnote{Note that this assumption is typically also made when individual country end-user prices are used (e.g. Griffin and Schulman 2005; Dargay and Gately 2010), which is more controversial. Notice also that, in contrast to most existing panel studies on oil demand, we do not have the US in our sample. Since the US has a share in global oil consumption of 27% over the sample period, endogeneity problems are more likely for panels that include the US.} A similar reasoning can be applied to the use of the US dollar effective exchange rate. When shifts in oil consumption of an individual country affect the bilateral US dollar exchange rate on impact, and this country has a large weight in the US dollar effective exchange rate index, the estimated elasticity could
be biased.\textsuperscript{21} Remark that, if such a bias is present, the true exchange rate elasticity of oil demand is probably even larger than the one we have reported above. Specifically, if an increase in a country’s oil demand raises its demand for US dollars and leads to an appreciation of the dollar, the estimated (negative) elasticity will decline.

To account for possible endogeneity between the demand for oil in individual (non-US dollar) countries, the global oil price, and US dollar exchange rate, we have re-estimated the baseline panel error correction oil demand model with instrumental variables (IV) as another robustness check. In particular, we instrument the first differences of the global price of crude oil and the US dollar real effective exchange rate in equation (1.2) by the first difference and the level of the US federal funds rate.

As shown in the left panel of Table 1.5, the results of the IV estimations confirm a relatively strong impact of the US dollar effective exchange rate on oil demand in the three samples but the effect is confined to the long-run coefficients. The differences in magnitudes relative to the global price coefficient even increase for the long-run coefficients, relative to the benchmark results reported in section 3.1. We obviously have to be careful when interpreting the magnitudes of the coefficients, given the loss of power of two-stage regressions with instrumental variables. This is reflected in the relatively large standard errors for the short-run price and exchange rate elasticities.

As a final robustness check, we have estimated a specification with bilateral real exchange rates instead of the US dollar real effective exchange rate. While the effective US dollar exchange rate reflects changes in the overall value of the US dollar, i.e. the currency unit which matters for global oil market dynamics, bilateral exchange rates capture more of the effects on the local oil price that consumers have to pay in each individual country. We again use IV estimations.\textsuperscript{22} The right panel of Table 1.5 shows the results. In general, the estimated coefficients turn out to be different from the benchmark results, which is probably due to the loss of power of the two-stage estimation procedure, and the difficulty

\textsuperscript{21}Japan has the largest weight in the US dollar effective exchange rate for our sample of countries, notably 18 percent since 1990.

\textsuperscript{22}Potential endogeneity problems between the bilateral US dollar exchange rate and the country-specific demand for oil are more likely than for the specification with the effective US dollar exchange rate. We apply the same instruments for the bilateral exchange rates, but use the lagged level of the real effective US dollar exchange rate instead of the lagged level of the bilateral exchange rate to instrument the oil price, given the common nature of the oil price variable. Notice that the sample size for the different groups of countries is now smaller due to the non-availability of bilateral exchange rate and/or domestic CPI data for some countries over the time period under consideration (see Appendix A).
to find proper instruments. In particular, the estimated global oil price elasticity becomes insignificant for the non-OECD group, and even significantly positive for the total sample of countries in the short run. The exchange rate elasticity estimates are, in contrast, still (significantly) negative in the short run. The magnitudes are however large, in particular for the non-OECD and total sample, while the standard errors increase considerably, which points to an inefficient estimator.

4 Pass-through of USD exchange rate to oil product prices

A striking result is that we consistently find that shifts in the US dollar exchange rate have a stronger impact on oil consumption in non-US dollar oil-importing countries than changes in the global price of crude oil expressed in US dollar. In this section, we analyze this in more detail. More specifically, we examine whether differences in the pass-through of global crude oil prices and the US dollar exchange rate to domestic end-user prices can explain the difference. The inertia of domestic prices of internationally traded goods to exchange rate changes is well-documented in international economics (e.g. Engel 2003; Goldberg and Hellerstein 2013) and might be different for shifts in global crude oil prices. Note that the analysis in this section is restricted to a confined group of 20 OECD countries, and only starts in 1978 due to the availability of end-user oil price data (see appendix). The pass-through analysis thus covers a substantially reduced sample in terms of the number of countries and time observations compared to the entire sample in the baseline oil demand model, but should nevertheless be instructive. The data for the G-7 oil-importing countries that are included in the analysis are shown in Figure 1.2.

In line with the existing literature on the pass-through of changes in the exchange rate to domestic good prices (e.g. Bussière 2013), we estimate a simple dynamic linear oil product price equation for the panel of 20 OECD countries over the sample period 1978-2008 of the following form:

\[ \Delta P_{it}^{\text{end}} = \alpha_0 + \alpha_1 \Delta \text{rac}_t + \alpha_2 \Delta \text{rer}_t + \alpha_3 \Delta P_{i,t-1}^{\text{end}} + \varepsilon_{it} \]  

where \( \Delta \text{rac}_t \) and \( \Delta \text{rer}_t \) are again the global price of crude oil expressed in US dollar and the real effective US dollar exchange rate, while \( \Delta P_{it}^{\text{end}} \) represents the domestic end-user oil products prices in local currency of country \( i \). These prices include taxes and are a weighted average of 4 product groups, i.e. motor gasoline, gas/diesel oil, light fuel oil

16
and residual oil. The results are shown in Table 1.6. According to Wald and CD tests, the preferred estimator is the MG estimator adjusted for cross-sectional dependence. The results reveal that both the pass-through of the US dollar effective exchange rate and the global crude oil price to domestic oil product prices is incomplete (i.e. less than proportional). The incomplete pass-through is interesting by itself, because it suggests for instance that studies which use the global price of crude oil to estimate the oil price elasticity of oil demand probably underestimate the true elasticity.

More importantly in the context of the present study, however, is the appreciable difference in the magnitudes of the $\alpha_1$ and $\alpha_2$ coefficients. In particular, the magnitude of the exchange rate pass-through is about twice as large as the global crude oil price pass-through to end-user product prices. The difference between both coefficients is statistically significant according to a Wald test. This finding suggests that a different pass-through of US dollar and global crude oil price fluctuations to end-user prices could be an explanation for the different impact on oil demand that we have found.

A stronger pass-through of the US dollar exchange rate to end-user prices relative to changes in the global price of crude oil, however, is not necessarily the only explanation for the larger exchange rate elasticity of oil demand. On top of this, the exchange rate could also affect other conditions that have an impact on oil demand (e.g. by also affecting the cost of borrowing). To further examine this, Table 1.7 presents the estimation results of the benchmark specification, where we have replaced the global price of crude oil expressed in US dollars by the domestic end-user product prices expressed in local currency. The results reveal that the oil price elasticity indeed increases when local prices are used, a finding which is consistent with Dargay and Gately (2010), and van Benthem and Romani (2009). Moreover, the exchange rate coefficient is now never statistically significant anymore, which indicates that there is no additional effect of the US dollar exchange rate on the demand for oil once the pass-through to end-user prices is incorporated in the oil price variable. In other words, we can conclude that differences in the pass-through to local oil product prices appear to be the key reason for the stronger effect of shifts in the US

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23 Since local oil prices could be affected by shifts in local oil demand, these results need to be taken with more than the usual degree of caution. Notice also that the inclusion of the exchange rate does not seem to absorb some of the true price and income effects. Specifically, as can be seen in Table 1.7, the price and income coefficients are very similar when we estimate the model without the US dollar exchange rate. Only the long-run price elasticity coefficient of the MG_Ft seems to be affected, but also the standard errors are quite large.

24 When we re-estimate this specification by converting the local oil product prices to US dollar, the
dollar exchange rate on oil demand relative to changes in global crude oil prices.

5 Economic relevance

In this section, we perform a back-of-the-envelope calculation in order to assess the relevance of US dollar exchange rate fluctuations for global oil market dynamics. Given its simplicity, the exact numbers should be interpreted with caution. The calculation does, for instance, not take into account endogenous dynamics. It should nevertheless give an idea about the importance of the US dollar for the oil market.

First, we assume that the estimated exchange rate elasticity for the total sample (−0.088) is representative for global (non-US) oil demand. Given the fact that the US represents on average 27 percent of world oil demand over the sample, this implies that a 1 percent appreciation of the US dollar leads to a decline in global oil demand by 

\[-0.088 \times (1 - 0.27) = -0.064\] percent. Based on this number, and the price elasticity estimate, consider the following simplified short-run oil demand function for the global oil market:

\[\Delta q_{oil} = -0.039\Delta p_{oil} - 0.064\Delta rer_{US}\]

where \(q_{oil}\), \(p_{oil}\) and \(rer_{US}\) are respectively global crude oil demand, the global real price of crude oil and the real effective US dollar exchange rate. Furthermore, according to Kilian and Murphy’s (2012) reading of the literature, the upper bound on the short-run price elasticity of oil supply is 0.025, which gives us the following simplified short-run global crude oil supply function:

\[\Delta q_{oil} < 0.025\Delta p_{oil}\]

Solving this model delivers the following effects of a shift in the US dollar exchange rate on oil prices and production:

\[|\Delta p_{oil}| > 1.004|\Delta rer_{US}|\]
\[|\Delta q_{oil}| > 0.025|\Delta rer_{US}|\]

As a benchmark, the monthly average of \(|\Delta rer_{US}|\) in the data is for instance 1.16 percent. According to our simple back-of-the-envelope calculation, this corresponds to an exchange rate coefficient becomes again significant, while the magnitude of the oil price elasticity is hardly affected. These results are available upon request.
average shift in global oil prices by 1.17 percent. Given the fact that the monthly average of $|\Delta p_{oil}|$ in the data is 4.76 percent, the relevance of US dollar exchange rate fluctuations for global oil price dynamics is considerable. Due to the very low oil supply elasticity, this is less the case for oil production. In particular, the monthly average of $|\Delta q_{oil}|$ in the data is 1.08 percent, whereas exchange rate fluctuations could only explain about 0.03 percent according to our simple calculations.

Average elasticities are, however, not necessarily representative for the global oil market, which is essentially a weighted average of all individual countries in the world. Therefore, we have also calculated weighted MG estimates of the panel error correction oil demand model, where the weights of the country-specific coefficients are determined by the share of the respective country in the total oil consumption over the sample. Accordingly, countries with a larger share in global oil demand have more weight such that the resulting MG estimates better represent global elasticities.\footnote{Our sample represents 59 percent of non-US global oil demand.} The short-run price and exchange rate coefficients that result from this exercise are respectively $-0.045$ and $-0.122$. These coefficients in turn result in an impact of US dollar effective exchange rate fluctuations of more than 1.48 percent on world oil price dynamics and 0.03 percent on global oil production using the above described procedure. In sum, our back-of-the-envelope calculations demonstrate that the effective US dollar exchange rate appears to be very important for fluctuations in global crude oil prices through its effect on the demand for oil. The effect on oil production on the other hand is very limited due to the very low short-run price elasticity of oil supply.

6 Conclusions

In this paper, we have examined the role of the US dollar exchange rate for the demand for oil in non-US dollar regions by using recent advances in panel data estimation techniques. In particular, we have estimated a panel error correction oil demand model allowing for cross-country heterogeneity in the slope coefficients, and taking into account cross-country common unobserved variables. The results show that an appreciation of the US dollar exchange rate robustly leads to a decline in the demand for oil in countries that do not use the US dollar for local transactions, which supports the premise of a significant exchange rate channel underlying oil demand dynamics. Strikingly, a 1 percent shift in the real US
dollar exchange rate seems to have a much stronger effect on oil demand than a 1 percent shift in the global real crude oil price determined in US dollar. A more detailed analysis of the effect of changes in the global crude oil price and the US dollar exchange rate on country-specific end-user prices of oil products suggest that the difference is the result of a much stronger pass-through of exchange rate fluctuations to end-user prices. The reason for the stronger pass-through of the US dollar effective exchange rate is beyond the scope of this paper, but could be a promising avenue for future research. A potential avenue is the lower volatility of the US dollar exchange rate compared to the global crude oil price.

A back-of-the-envelope calculation furthermore suggests that shifts in the US dollar exchange rate are economically important for global (US dollar) crude oil price fluctuations due to its influence on global oil demand. It is thus recommended to include the US dollar exchange rate in the analysis of global oil market dynamics in the spirit of Kilian (2009), Peersman and Van Robays (2009) and Juvenal and Petrella (2012).
Data Appendix

Data sources:

- Total oil demand (1000 barrels per day): International Energy Agency (IEA), Oil Information database
- Total midyear population (number of persons): US Census Bureau, International database
- Global crude oil price (US dollars per barrel): Energy Information Administration (EIA), Refiner acquisition cost of imported crude oil
- Real gross domestic product per capita.: Penn World Tables 7.0, PPP Converted GDP Per Capita (Chain Series), 2005 constant prices
- Real US dollar effective exchange rate: BIS, real effective exchange rate index (CPI-based), Narrow Index (2010=100)
- Monthly Crude oil and NGL production for Figure 1.1 (1000 barrels per day): IEA, Oil Information database
- Individual country nominal exchange rates [ER] (national currency unit to US $, period average) : IMF, IFS database
- Consumer prices [CPI] (indices, 2005=100): IMF, IFS database
- End-user oil products prices [penduser]: IEA, Energy Prices and Taxes database
- US Federal Funds rate: Datastream, US Federal Funds effective rate (code: FRFEDFD?)

Construction variables:  
- Total oil demand per capita [DEM] (barrels per day, per 1000 persons) = Total Oil demand/ Total population*1000
- Real global crude oil price [POIL]: Global nominal crude oil price/ CPIust*100
- Real exchanges rates [bRER] = ERit * CPIus,t / CPIit, index 2005=100
-> Real end-user oil products price indexes in national currencies [penduser]: quantity weighted real end-user price index (2005=100) of oil products based on Griffin and Schulman (2005), i.e. for 4 product groups: residual oil, light fuel oil, motor gasoline and gas/diesel oil weighted based on their respective importance in aggregated consumption.

Sample coverage: The dataset is balanced for 65 oil-importing countries over the sample period 1971-2008. All variables are converted to natural logarithms, such that the models are of the constant elasticity form.

OECD sample (23 countries): Australia, Austria, Belgium, Denmark, Finland, France, Germany, Greece, Hungary, Iceland, Ireland, Italy, Japan, South Korea, Luxembourg, Netherlands, New Zealand, Poland, Portugal, Spain, Sweden, Switzerland and Turkey

Non-OECD sample (42 countries): Bangladesh, Benin, Bolivia, Brazil, Bulgaria, Chile, China, Costa Rica, Cote d'Ivoire, Cyprus, Dominican Republic, El Salvador, Ethiopia, Ghana, Guatemala, Haiti, Honduras, Hong Kong, India, Israel, Jamaica, Jordan, Kenya, Malta, Morocco, Mozambique, Nicaragua, Pakistan, Paraguay, Peru, Philippines, Romania, Senegal, Singapore, South Africa, Sri Lanka, Sudan, Tanzania United Republic, Thailand, Uruguay, Zambia and Zimbabwe

The following countries have been excluded from the analysis because of being a net oil-exporting country, i.e. countries for which the production of crude oil has been larger than total oil demand for at least 25 years: Canada, Mexico, Norway, United Kingdom, Algeria, Angola, Argentina, Bahrain, Cameroon, Colombia, Congo, Congo Democratic Republic, Ecuador, Egypt, Gabon, Indonesia, Iran, Iraq, Kuwait, Malaysia, Nigeria, Oman, Qatar, Saudi Arabia, Syrian Arab Republic, Tunisia, United Arab Emirates and Venezuela

When the model includes the bilateral real US dollar exchange rate instead of the real effective US dollar exchange rate index, the total sample reduces to 44 countries (20 OECD and 24 non-OECD countries) due to missing data for the bilateral nominal exchange rates and/or consumer price indices for the entire time period under consideration.

-> Missing OECD: Hungary, Poland, Turkey

-> Missing non-OECD: Bangladesh, Benin, Bolivia, Brazil, Bulgaria, Chile, China, Ghana, Hong Kong, Israel, Mozambique, Nicaragua, Peru, Romania, Sudan, Uruguay,
Zambia, Zimbabwe

When the model includes end-user oil product prices instead of the real global crude oil price, the total sample reduces to 20 OECD countries due to missing data on end-user oil products price indices for the entire time period under consideration (1978-2008).

-> Missing OECD: Australia, Iceland and Turkey
Bibliography


Figure 1.1: evolutions in oil consumption in G-7 countries in sample and in total OECD and non-OECD aggregates versus the evolution of real effective US dollar exchange rate

Note: the left axis’ units refer the percentage deviation from trend from total oil demand per capita (barrels per day, per 1000 persons), the right axis refers to the real effective US dollar exchange rate (RER) which is an index equal to 100 in base year 2005.

Sources data: total oil demand: IEA / population individual countries: US Census Bureau (international database) / population OECD and non-OECD country aggregates: OECD (population database) / RER: BIS (narrow index).
Note: All series are indexed with base year 2005 equal to 100. The individual country series refer to quantity weighted real end-user price indexes of oil products based on Griffin and Schulman (2005), i.e. for 4 product groups: residual oil, light fuel oil, motor gasoline and gas/diesel oil weighted based on their respective importance in aggregated consumption, the RER series refers to the real effective US dollar exchange rate and the OILP series to the global real crude oil price.

Sources data: end-user oil products prices for individual countries: International Energy Agency (Energy Prices and Taxes database) / OILP: Energy Information Administration website (refiner acquisition cost of imported crude oil) / RER: BIS (narrow index).
### Table 1.1: Results PANIC tests, total sample

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#### Intercept only model

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<td>-0.93</td>
<td>-0.18</td>
</tr>
<tr>
<td>MQf</td>
<td>-1.58</td>
<td>-0.64</td>
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</table>

#### Linear trend model

<table>
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<th>DEMit</th>
<th>GDPit</th>
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<tbody>
<tr>
<td>MW</td>
<td>163.70**</td>
<td>117.91</td>
</tr>
<tr>
<td>Choi</td>
<td>2.09**</td>
<td>-0.75</td>
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<tr>
<td>ADFf</td>
<td>-</td>
<td>-</td>
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<tr>
<td>MQc</td>
<td>-6.09</td>
<td>-4.8</td>
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<tr>
<td>MQf</td>
<td>-7.12</td>
<td>-6.43</td>
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#### ADF test common observed variables

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<td>QLipt</td>
<td>-1.25</td>
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<tr>
<td>RERT</td>
<td>-3.26**</td>
<td>-3.72**</td>
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### Table 1.2: Results panel cointegration tests, total sample

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<tr>
<th></th>
<th>Model 1</th>
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<th>Model 3</th>
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<tbody>
<tr>
<td>$t_{pi}$</td>
<td>-1.47</td>
<td>-5.66***</td>
<td>-6.11***</td>
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<tr>
<td>$\omega_{pi}$</td>
<td>24.64***</td>
<td>29.12***</td>
<td>34.20***</td>
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</table>

GUE test [Gengenbach, Urbain and Westerlund, 2008] pooled tests for no error correction

Number of lags determined by Akaike Information Criterion (AIC)

$\omega_{pi}^*$ = average of truncated version of the individual t-test statistics of no error correction

Model 1 = model with no deterministic terms

Model 2 = model with unrestricted constant

Model 3 = model with unrestricted constant and trend

**PANIC test on CupBC residuals:**

First step: Continuously updated bias-correcting (CupBC) estimator (Bai et al, 2003) on static long-run specification

Second step: PANIC test (Bai and Ng, 2004) on idiosyncratic part of residuals of first-step regression

---

**PANIC test (Bai and Ng, 2004):**

Number of common factor determined by IC1, IC2, and BIC3 criteria (Bai and Ng, 2002)

Number of lags in both model specifications: determined by Bai and Ng (2004) rule: $4/\min(N,T)$

MW = Maddala and Wu (1999) pooled unit root test statistic on idiosyncratic term

Choi = Choi (2001) pooled unit root test statistic on idiosyncratic term

ADP = Augmented Dickey Fuller (ADF) test (Dickey and Fuller, 1979) on estimated common factor if $r=1$

MQc/MQf = Modified variants of Stock and Watson’s (1988) Qf and Qc statistics to determine the number of factors spanning the non-stationary space of the common term

Augmented Dickey Fuller (ADF) test on observed common factors

Lag order determined by Schwarz Bayesian Criterion (SBC)
### Table 1.3: baseline model

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<tr>
<th></th>
<th>OECD</th>
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<th>non-OECD</th>
<th></th>
<th>TOTAL</th>
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<tbody>
<tr>
<td></td>
<td>FE</td>
<td>MG</td>
<td>MG_Ft</td>
<td>FE</td>
<td>MG</td>
<td>MG_Ft</td>
</tr>
<tr>
<td>gdp</td>
<td>0.759**</td>
<td>0.667**</td>
<td>0.686***</td>
<td>0.494**</td>
<td>0.500**</td>
<td>0.532***</td>
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<tr>
<td></td>
<td>(0.084)</td>
<td>(0.080)</td>
<td>(0.077)</td>
<td>(0.102)</td>
<td>(0.085)</td>
<td>(0.093)</td>
</tr>
<tr>
<td>oilp</td>
<td>-0.047**</td>
<td>-0.052**</td>
<td>-0.054***</td>
<td>-0.029**</td>
<td>-0.025*</td>
<td>-0.027**</td>
</tr>
<tr>
<td></td>
<td>(0.007)</td>
<td>(0.008)</td>
<td>(0.008)</td>
<td>(0.010)</td>
<td>(0.013)</td>
<td>(0.013)</td>
</tr>
<tr>
<td>rer</td>
<td>-0.169***</td>
<td>-0.150**</td>
<td>-0.192***</td>
<td>-0.133**</td>
<td>-0.060</td>
<td>-0.057</td>
</tr>
<tr>
<td></td>
<td>(0.030)</td>
<td>(0.033)</td>
<td>(0.035)</td>
<td>(0.051)</td>
<td>(0.056)</td>
<td>(0.056)</td>
</tr>
<tr>
<td>trend</td>
<td>-0.002***</td>
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<td>-0.001</td>
<td>-0.000</td>
<td>-0.001</td>
<td>-0.001</td>
</tr>
<tr>
<td></td>
<td>(0.000)</td>
<td>(0.002)</td>
<td>(0.001)</td>
<td>(0.000)</td>
<td>(0.002)</td>
<td>(0.002)</td>
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</table>

**long-run coefficients**

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<tr>
<th></th>
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<th>MG_Ft</th>
<th>FE</th>
<th>MG</th>
<th>MG_Ft</th>
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</thead>
<tbody>
<tr>
<td>EC</td>
<td>-0.074***</td>
<td>-0.244***</td>
<td>-0.202***</td>
<td>-0.110***</td>
<td>-0.429**</td>
<td>-0.391***</td>
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<tr>
<td></td>
<td>(0.010)</td>
<td>(0.037)</td>
<td>(0.045)</td>
<td>(0.030)</td>
<td>(0.039)</td>
<td>(0.038)</td>
</tr>
<tr>
<td>gdp</td>
<td>1.108***</td>
<td>0.915**</td>
<td>0.520**</td>
<td>0.732***</td>
<td>1.013**</td>
<td>1.064***</td>
</tr>
<tr>
<td></td>
<td>(0.010)</td>
<td>(0.416)</td>
<td>(0.219)</td>
<td>(0.018)</td>
<td>(0.157)</td>
<td>(0.177)</td>
</tr>
<tr>
<td>oilp</td>
<td>-0.383***</td>
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<td>-0.286***</td>
<td>-0.115**</td>
<td>-0.151***</td>
</tr>
<tr>
<td></td>
<td>(0.005)</td>
<td>(0.023)</td>
<td>(0.041)</td>
<td>(0.011)</td>
<td>(0.033)</td>
<td>(0.037)</td>
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<tr>
<td>rer</td>
<td>-0.689***</td>
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<td>-0.308***</td>
<td>-0.322**</td>
<td>-0.343***</td>
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<tr>
<td></td>
<td>(0.022)</td>
<td>(0.096)</td>
<td>(0.102)</td>
<td>(0.031)</td>
<td>(0.089)</td>
<td>(0.088)</td>
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</table>

**Swamy’s Wald test**

| statistic | 318.57*** | 468.00*** | 833.45*** |

**Cross section Dependence test**

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<tr>
<th>statistic</th>
<th>$\rho$</th>
<th>10.93***</th>
<th>10.29***</th>
<th>5.67***</th>
<th>4.84***</th>
<th>5.38***</th>
<th>4.59***</th>
<th>13.01***</th>
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<th>11.01***</th>
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<tbody>
<tr>
<td></td>
<td>$\rho$</td>
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<td>0.059</td>
<td>0.027</td>
<td>0.030</td>
<td>0.026</td>
<td>0.047</td>
<td>0.042</td>
<td>0.040</td>
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</tbody>
</table>

FE= Fixed Effects estimator
MG= Mean Group estimator
MG_Ft= Mean Group estimator adjusted for cross-sectional dependence
$p$ = average pair-wise correlation coefficient of residuals

***/**/*: respectively refers to significance at the 1/5/10 % level
standard errors are listed in brackets (robust s.e. for FE)

The empirical analysis was carried out in Stata 12, and we employed the user-written Stata routines xtd and xtmg written by Markus Eberhardt (Eberhardt, 2012). An outlier-robust weighting procedure is used to construct the MG estimates, following Bond, Lelebicioglu and Schiantarelli (2010).
## Table 1.4: price decomposition

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<tbody>
<tr>
<td></td>
<td>FE</td>
<td>MG</td>
<td>MG_Ft</td>
<td>FE</td>
<td>MG</td>
<td>MG_Ft</td>
<td>FE</td>
<td>MG</td>
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<td>FE</td>
</tr>
<tr>
<td>gdp</td>
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<td>0.408***</td>
<td>0.467***</td>
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<td>(0.078)</td>
<td>(0.083)</td>
<td>(0.088)</td>
<td>(0.060)</td>
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<tr>
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<td>-0.109***</td>
<td>-0.108***</td>
<td>-0.096***</td>
<td>-0.093***</td>
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<td>-0.101***</td>
<td>-0.098***</td>
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<td>(0.018)</td>
<td>(0.019)</td>
<td>(0.019)</td>
<td>(0.012)</td>
<td>(0.012)</td>
<td>(0.012)</td>
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<tr>
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<td>-0.042***</td>
<td>-0.062*</td>
<td>-0.006</td>
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<td>-0.022**</td>
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<td>(0.010)</td>
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<tr>
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<td>0.002</td>
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<td>(0.002)</td>
<td>(0.002)</td>
<td>(0.004)</td>
<td>(0.003)</td>
<td>(0.004)</td>
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### Long-run coefficients

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<td>(0.043)</td>
<td>(0.041)</td>
<td>(0.024)</td>
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<td>(0.033)</td>
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<tr>
<td>gdp</td>
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<td>0.766***</td>
<td>0.825***</td>
<td>0.845***</td>
<td>0.731***</td>
<td>0.766***</td>
<td>0.819***</td>
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<td>(0.012)</td>
<td>(0.165)</td>
<td>(0.245)</td>
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<td>(0.133)</td>
<td>(0.015)</td>
<td>(0.114)</td>
<td>(0.109)</td>
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<tr>
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<td>-0.245***</td>
<td>-0.223***</td>
<td>-0.840***</td>
<td>-0.245***</td>
<td>-0.233***</td>
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<tr>
<td></td>
<td>(0.011)</td>
<td>(0.045)</td>
<td>(0.043)</td>
<td>(0.025)</td>
<td>(0.041)</td>
<td>(0.039)</td>
<td>(0.018)</td>
<td>(0.030)</td>
<td>(0.031)</td>
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<tr>
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<td>(0.055)</td>
<td>(0.060)</td>
<td>(0.012)</td>
<td>(0.034)</td>
<td>(0.036)</td>
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<tr>
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<td>0.036**</td>
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<td>(0.039)</td>
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<td>(0.022)</td>
<td>(0.061)</td>
<td>(0.063)</td>
<td>(0.014)</td>
<td>(0.039)</td>
<td>(0.042)</td>
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<tr>
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<td>-0.713***</td>
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<td>-0.281***</td>
<td>-0.302***</td>
<td>-0.563***</td>
<td>-0.329***</td>
<td>-0.329***</td>
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<tr>
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<td>(0.020)</td>
<td>(0.076)</td>
<td>(0.108)</td>
<td>(0.031)</td>
<td>(0.065)</td>
<td>(0.057)</td>
<td>(0.012)</td>
<td>(0.049)</td>
<td>(0.044)</td>
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### Swamy's Wald test

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<tr>
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<th>565.28***</th>
<th>964.91***</th>
<th>1586.41***</th>
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<td></td>
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<tr>
<td>statistic</td>
<td>4.80***</td>
<td>9.33***</td>
<td>4.70***</td>
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<tr>
<td>p</td>
<td>0.050</td>
<td>0.098</td>
<td>0.049</td>
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</table>
\begin{table}
\centering
\begin{tabular}{|c|c|c|c|c|c|c|}
\hline
 & \textbf{Real effective US $ exchange rate} & & \textbf{Real bilateral US $ exchange rates} & & \\
 & \textbf{OECD (N=23)} & \textbf{non-OECD (N=42)} & \textbf{TOTAL (N=65)} & \textbf{OECD (N=20)} & \textbf{non-OECD (N=24)} & \textbf{TOTAL (N=44)} \\
\hline
\textbf{gdp} & 0.731*** & 0.537*** & 0.629*** & 0.567*** & 0.184 & -0.127 \\
 & (0.149) & (0.124) & (0.080) & (0.086) & (0.252) & (0.142) \\
\hline
\textbf{oilp} & -0.141** & -0.076 & -0.107*** & -0.053** & -0.014 & 0.077*** \\
 & (0.070) & (0.066) & (0.039) & (0.026) & (0.036) & (0.027) \\
\hline
\textbf{rer / brer} & -0.086 & 0.112 & -0.142 & -0.181*** & -1.408** & -1.309*** \\
 & (0.164) & (0.129) & (0.100) & (0.064) & (0.579) & (0.194) \\
\hline
\textbf{trend} & -0.001 & -0.001 & -0.003** & -0.001 & 0.010** & 0.012*** \\
 & (0.002) & (0.002) & (0.001) & (0.002) & (0.005) & (0.002) \\
\hline
\hline
\multicolumn{7}{|c|}{\textbf{long-run coefficients}} \\
\hline
\textbf{EC} & -0.291*** & -0.352*** & -0.288*** & -0.258*** & -0.501*** & -0.452*** \\
 & (0.053) & (0.035) & (0.025) & (0.049) & (0.071) & (0.033) \\
\hline
\textbf{gdp} & 0.832*** & 0.260 & 0.443*** & 0.103 & 0.624** & 0.718*** \\
 & (0.226) & (0.180) & (0.146) & (0.130) & (0.226) & (0.242) \\
\hline
\textbf{oilp} & -0.027 & -0.050** & -0.054*** & -0.013** & -0.021 & 0.013 \\
 & (0.020) & (0.022) & (0.015) & (0.005) & (0.041) & (0.020) \\
\hline
\textbf{rer / brer} & -0.289** & -0.622*** & -0.429*** & -0.004 & -0.008 & 0.029 \\
 & (0.140) & (0.117) & (0.079) & (0.008) & (0.071) & (0.044) \\
\hline
\hline
\textbf{Cross section Dependence test} & & & & & & \\
\hline
\textbf{statistic} & 4.46*** & 7.86*** & 17.62*** & 5.98*** & 5.62*** & 17.31*** \\
\hline
\textbf{\(\rho\)} & 0.046 & 0.044 & 0.064 & 0.071 & 0.056 & 0.093 \\
\hline
\end{tabular}
\end{table}

IV=dFFR\(_{it}\), FFR\(_{it}\) & IV=dFFR\(_{it}\), FFR\(_{it}\) for \(d\text{rer}_{it}\)/dFFR\(_{it}\), FFR\(_{it}\), REER\(_{it}\) for \(d\text{rac}_{it}\)
<table>
<thead>
<tr>
<th></th>
<th>OECD (N=20)</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>FE</td>
</tr>
<tr>
<td>Δpenduser</td>
<td></td>
</tr>
<tr>
<td>Δrac</td>
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</tr>
<tr>
<td></td>
<td>(0.024)</td>
</tr>
<tr>
<td>Δrer</td>
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CHAPTER 2

Traditional export determinants and export dynamics since the Great Trade Collapse: a cross-country analysis.

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Abstract

By means of a panel model of goods exports for 16 OECD economies, we quantify advanced economies' export performance since the "Great Trade Collapse" (GTC) based on traditional export determinants. We include a variable measuring shifts in the sectoral composition of world trade and split the real exchange rates up into its constituent parts to account for a different response to unit labour costs and the nominal exchange rate. We find that, on average, a pre-crisis model explains aggregate exports since the GTC well once one controls for unobserved common factors. We do find substantial cross-country variation in export performance based on the average coefficients.

JEL classification: C23, F14, F17

Keywords: International trade, forecasting, cross-country panel

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1 Introduction

In the immediate aftermath of the financial crisis, the sharp falls in output across developed economies were accompanied by even stronger falls in international trade. This fall in trade, often termed "The Great Trade Collapse" (GTC) was by far the largest drop in global trade in history, as global trade plunged by 15% year-on-year in late 2008 and early 2009. In addition, the GTC was characterised by an extremely high level of synchronization with falls in exports across all advanced economies and nearly all industries (Baldwin, 2009; IMF, 2013).

The dynamics of world trade since the GTC reheated the long-standing debate on the relationship between trade and macroeconomic dynamics. The consensus is that the fall in world output and industrial production that occurred at the same time played a key part in the GTC. But some authors have argued that other factors outside of traditional trade models contributed to the sharp fall in trade such as impaired trade finance (Chor and Manova, 2012; JaeBin, Amiti and Weinstein, 2011) or heightened uncertainty (Novy and Taylor, 2014). Others have argued that the GTC heralded an era of lower world trade relative to GDP - either because aspects of globalisation such as offshoring may have run their course (Krugman, 2013) or because of a rise in hidden protectionism (Davies, 2013). Whilst the former imply shorter-run shocks that may reverse over time as the global economy recovered, the latter suggest a more permanent hit to trade due to a structural break in the relationship between economic activity and trade. At the same time, export growth following the GTC has been argued to be further constrained by the fact that global growth remained relatively subdued. Some policymakers have highlighted the difficulties posed by weaker external conditions for some economies seeking to rebalance away from domestic demand towards exports.

Our focus in this paper is to analyse what has happened to exports of advanced economies since the GTC and to gauge how consistent they were with pre-crisis export models, rather than to uncover the cause of the GTC per se. In particular, we construct a forecast benchmark for the actual exports of a panel of advanced economies to analyse their export evolutions since the onset of the GTC. The capability of a pre-crisis standard export relationship to account for the evolutions of exports during and following the GTC is of great interest given its important policy implications. If the GTC marks structural changes in the relationship between trade performance and the traditional macroeconomic
determinants, counterfactual export analyses based on traditional export determinants can provide misleading policy guidance.

In this work, we first extend the standard export equation based on the imperfect substitutes model of international trade (Goldstein and Kahn, 1985) to incorporate the effects of shifts in the sectoral composition of global trade and the role of imported inputs and local nontraded costs on the effect of the real exchange rate. We analyse this model for goods exports in a panel error correction model (ECM) framework for 16 advanced economies. Whereas recent empirical work using ECMs to model exports estimated either a system of country specific equations or focused on aggregate exports\(^1\), we instead adopt a panel approach using the Common Correlated Effects (CCE) estimator of Pesaran (2006) to control for the possible existence of unobserved common factors.

We develop a novel measure of sectoral shifts in world trade and examine its effects next to the traditional price and income determinants of exports as compositional effects have been found to be important for trade in the long run (Mayer, 2010) and in the short run as well (Levchenko, Lewis and Tesar, 2010). The measure allows us to control for compositional demand effects on export flows. Note that this is distinct from a decomposition of aggregate demand into its expenditure components according to their trade intensity as employed in Bussière \textit{et al} (2013), because it divides the data up along sectoral lines.\(^2\) We find the variable to have only a significant lagged impact on export in the short run.

We further split up the real exchange rate (our preferred measure of relative prices) into its two constituent parts, the nominal exchange rate and relative unit labour costs to differentiate between the effects of the nominal exchange rate and relative cost changes which might be different because of the importance of nontraded costs and because marginal producer costs strongly co-move with exchange rates due to imported inputs (Goldberg and Hellerstein, 2008). Both factors can limit the responsiveness of exports to nominal exchange rates because they affect the pricing behavior of producers. A depreciation of

\(^1\)Ca’Zorzi and Schnatz (2007) estimate an ECM for aggregate euro area exports, with a focus on assessing the best measure of competitiveness from a forecasting perspective. di Mauro, Rüffer and Bunda (2008) estimate country specific export demand equations for France, Germany and Italy, as well as a pooled version for the three. Breuer and Klose (2013) estimate individual export demand equations for 9 euro area countries using the SURE methodology.

\(^2\)Next to the fact that we focus on compositional effects of global exports, the creation of an aggregate export content weighted demand measure similar to the import content weighted measure in Bussière \textit{et al} (2013) is also infeasible to construct given the lack of export content data.
the local currency renders imported inputs more expensive while nontraded costs are invariant to exchange rate changes. We find that the source of the real exchange rate shock matters for exports. The coefficients for both components of the real exchange rate differ significantly - the response to relative unit labour costs is around six times as large as to nominal exchange rates in the short run.

We next use the model estimated over a pre-crisis sample period to undertake a forecast exercise to explore the dynamics of trade since the GTC. This exercise allows to construct a benchmark to which the advanced economies’ exports since the GTC can be compared. First, we average the individual forecasts to yield an aggregate forecast. Comparing this forecast benchmark with the actual average outturn allows us to assess how the export flows since the GTC compare to the predictions of the pre-crisis model. Since the GTC, the dynamics of the out-of-sample forecast are highly similar to actual average exports. The out-of-sample forecast based on the pre-crisis model average coefficients can explain about 95 percent of the observed fall in exports between peak and through, i.e. between 2008Q2 and 2009Q2. Unobserved common factors, proxied by the cross section averages of all variables, constitute a major part of the predicted decline in advanced economies’ exports during the GTC.

Second, we examine the forecasts for each country conditional on the country-specific variables and the common factors to gauge each country’s performance against its peers. The forecasts in this case serve as uniform benchmarks to evaluate countries relative export performance based on the estimated average coefficients of the pre-crisis export model. The performance of the individual countries relative to their benchmark varies considerably across countries. From this, we can infer that advanced economies’ exports since the GTC relate to a varying degree to the average prediction based on the traditional demand and competitiveness determinants.

The remainder of the paper is organised as follows: section 2 reviews the relevant literature on the traditional export determinants, section 3 describes the specification of our model and outlines our data and estimation approach, section 4 presents the empirical results, section 5 the conditional forecasts and section 6 concludes.

3The forecasts are conditional on both the country-specific variables and the cross section averages of all variables in the model (which account for the unobserved common factors in the residual of the model).
2 Changing nature of international trade

This paper is part of the empirical literature on international economics that focuses on the estimation of trade elasticities and that dates back to the seminal work of Houthakker and Magee in 1969. The imperfect-substitutes model of international trade (Goldstein and Khan, 1985) offers the standard formulation of an export equation. We apply this theoretical framework to analyse the export performance of 16 OECD countries since the GTC.

The standard model is however more and more affected by the changing role of trade in the economy. First, as economies have become more opened to international trade by reducing trade barriers and tariffs, the overall volume of world trade has increased. Alongside traditional Ricardian channels, this evolution has increased the attractiveness of offshoring certain stages of the supply chain, often to lower cost producers (Strauß, 2002; Kleinert and Zorell, 2012; Yi, 2013). Offshoring (alternatively called vertical specialisation or outsourcing) results in an enlarged trade in intermediate goods between the offshoring firm and the foreign intermediate good producer and further boosts the volume of world trade (Kleinert and Zorell, 2012; Yi, 2013). Globalisation is likely to have increased the trade intensity of GDP, as lower trade barriers, transport costs and tariffs have led to a lengthening in supply chains.

This evolution of the nature of international trade has important repercussions for the standard trade equation as the income elasticity coefficients of exports ends up to be a combination of the traditional income effect and the effect of globalisation induced growth spurring trade. Whilst the former points to an import elasticity of no more than unity, the latter can generate a higher elasticity since trade is measured in gross and output in net terms, such that a dollar of extra output may be associated with more than a dollar of extra trade flows. The coefficient on GDP should thus not be interpreted as measuring the traditional income effect but as a combination of this income effect and the effect of increased globalisation.4

A separate channel is that globalisation is associated with increased international competition. As it is commonly assumed that the sensitivity of exports to cost changes depends

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4 Recently, the OECD and WTO (2012) have published a dataset of bilateral trade flows in gross value added (GVA). However, since this is only available for a single year (2009), it is unsuitable for the kind of dynamic panel analysis undertaken here.
on the degree of competition in the market (Carlin, Glyn and Van Reenen, 2001), also the elasticity of exports to cost changes can be altered. On the other hand, an enlarged import content of exports due to offshoring should lead to a decreased responsiveness to exchange rate changes (di Mauro et al, 2008).

Second, there may be an important role for compositional effects in international trade due to structural shifts in global demand (Mayer, 2010). Countries that specialise in the sectors in which world demand growth is concentrated or that are better able to respond to changes in the structure of international demand by adjusting their production accordingly, are likely to gain market share in global export markets. So even preference shifts that do not necessarily result in quantity effects at the aggregate level can be important for countries’ exports depending on which sectors are affected.

In this paper, we address both issues by refraining from imposing a unit coefficient on the income elasticity coefficient, by controlling for unobserved common factors and their possible correlation with the export determinants in the model and by introducing a proxy for the changing composition of trade for a given level of international trade.

3 Empirical methodology

3.1 Imperfect substitutes model

Our starting point is the standard export equation based on the imperfect substitutes model of international trade (Goldstein and Kahn, 1985).

\[ X_{i,t} = f(Y^*_{i,t}, E_{i,t}) \]  

(2.1)

Where \( X \) is real exports in domestic currency, \( Y^* \) is a country-specific trade-weighted external demand measure, and \( E \) is real effective exchange rates. Domestic exports are a function of foreign income, the domestic price of goods and the price of goods that compete with the reporting country’s goods in foreign markets. A rise in overseas output and a lower real exchange rate are expected to lead to an increase in exports.

\[ The \ CCE \ estimators \ allow \ to \ account \ for \ the \ influence \ of \ unobserved \ common \ factors \ by \ including \ the \ cross-sectional \ averages \ of \ the \ dependent \ and \ independent \ variables \ (see \ page \ 49). \]
We extend this basic setup along two different dimensions. First, we allow for changes in the sectoral composition of world trade to also play an important role in determining exports, and to do so in a way that may differ across countries. For example, lower cost emerging market economies (EMEs) may have displaced more established advanced economy producers in certain product types such as clothing and footwear. The effect of this shift in the sectoral composition of world trade will affect countries differently, depending on whether they specialise in the sectors most affected. On the other hand, sectoral shifts may also be beneficial. For example, whereas the group of rich developed countries’ demand tends to center on manufactured consumer goods, rapidly industrialising emerging economies’ demand is more focused on industrial raw materials, energy and food products (Mayer, 2010). Countries specialising in these products may benefit more from the growth in emerging economies than those who do not. Compositional effects however do not only relate to medium or long term structural changes, they can be important in the short run as well (Levchenko et al, 2010). Given the potential importance of compositional shifts in international trade, we add a variable \(C^*\) to the benchmark model to explore the effect of sectoral preferences for advanced economies’ exports, which we spell out in more detail in the next section.

Second, alongside the more standard approach of including a real exchange rate variable, we also explore the consequences of decomposing the real exchange rate into separate nominal exchange rate and domestic costs terms, i.e. \(E = \{S, U\}\), following Carlin et al (2001), Allard et al (2005), Breuer and Klose (2013) and Chen, Milesi-Ferretti and Tressel (2013). There are several reasons why the response of exports to a given real exchange rate shock may depend on which of the components is driving the change. If production of exported goods requires the use of imported inputs, then a given nominal depreciation may be partially offset by a rise in non-labour production costs, that would not occur if the real depreciation occurred because of an improvement in unit labour costs. Also the presence of nontraded local costs reduces the response of prices to nominal exchange rates (Goldberg and Hellerstein, 2008) and implies pricing-to-market (i.e. exchange rate changes are associated with markup variation). Alternatively, Obstfeld and Rogoff’s (2001) “exchange rate disconnect puzzle” highlights the fact that nominal exchange rates are far more volatile than fundamentals and that in the short run, the correlations between ex-

\footnote{For example, Giovanetti, Sanfilippo and Velucchi (2012) find evidence that Italy has been much more adversely affected by the rise of Chinese exports than Germany because Italy’s exports are in sectors with greater competition from China.}
ports and nominal exchange rate appear low. Given the larger persistence of relative costs compared to nominal exchange rates, exporters may react more swiftly to cost changes than to movements in nominal exchange rates.

Equation (2.1) is thus extended as follows:

\[ X_{i,t} = f \left( Y^*_i, S_{i,t}, U_{i,t}, C^*_{i,t} \right) \]  \hspace{1cm} (2.2)

where \( S \) refers to nominal exchange rates, \( U \) captures relative domestic costs and \( C^* \) is a measure of sectoral shocks.

### 3.2 Data

We estimate our equation over an unbalanced panel at the quarterly frequency, comprising 16 advanced economies\(^7\) between 1984Q1 and 2008Q2. A panel approach provides us with estimates of average elasticities over a comparable group of countries which can be used to examine individual countries’ export performances relative to their peers.

Our dependent variable, \( X \), is real export volumes in domestic currency for the goods sector. Our preferred measure of international price competitiveness, \( E \), is the IMF’s unit labour cost (ULC) based real effective exchange rate (REER) index. This captures relative unit labour costs versus competitors, using double weights to capture import and export competition in third markets.\(^8\) Export prices are also commonly used as a measure of international price competitiveness but they suffer from the drawback of being determined endogenously with export quantities. In addition, they may be largely influenced by pricing-to-market effects or other pricing behaviour, which are conceptually distinct from true cost competitiveness.

We prefer the ULC based REER to a consumer price index (CPI) based index because unit labour costs are likely to better reflect underlying cost shocks to producer prices. The CPI basket includes non-traded goods, regulated prices and services which may be a misleading indicator of traded goods prices. In addition, for many countries the CPI basket includes a substantial imported component, which can results in an understatement of the

\(^7\)The 16 countries are: Australia, Austria, Canada, Denmark, Finland, France, Germany, Italy, Japan, Netherlands, New Zealand, Portugal, Sweden, Switzerland, UK and the US.

\(^8\)For more details on the methodology, see Bayoumi, Lee and Jayanthi (2005).
competitiveness effects of a nominal exchange rate depreciation. By contrast, ULCs for the manufacturing sector\(^9\) give a broad indication of international price competitiveness as the manufacturing sector is representative for traded goods and labour costs represent a major component of total costs per unit of output. Moreover, by focusing on costs rather than prices, the competitiveness indicator is less subject to direct exchange rate effects on pricing behaviour. A ULC based REER is nevertheless by no means perfect as it also abstracts from indirect taxes and non-labour costs such as costs of raw materials or capital costs. We split the ULC based REER into its two subcomponents -relative ULCs and the nominal exchange rate. We do so by subtracting the equivalent (logged) nominal exchange rate series from its real counterpart to yield a measure of relative ULCs.

To measure the role of external demand conditions, we calculate trade-weighted world output growth \((\Delta Y^*)\), given by:

\[
\Delta Y^*_{i,t} = \sum_{p=1}^{P} \omega_{ipt} \Delta Y_{p,t} 
\]

(2.3)

where \(\omega_{ipt}\) is the weight of country \(i\)'s exports at time \(t\) going to partner \(p\), where the weights are given by a three-year moving average of the export shares calculated for 76 trading partners using bilateral trade flows.\(^{10}\) \(\Delta Y_{p,t}\) is real GDP growth in each trading partner. The growth rate of trade-weighted world output growth is subsequently used to construct the indexed \((2009Q1=100)\) level variable \(Y^*\).

We prefer this over an index of partners' import growth for a couple of reasons. Over the sample period, part of the rise in advanced economy imports reflects the supply-side shock of greater exports from emerging markets. The latter will either be orthogonal to demand for advanced economies -and hence a source of unwelcome noise- or even negatively correlated if they compete away advanced economies’ market share. In addition, because our left hand side variable covers advanced economies accounting for the bulk of world trade, trade-weighted world import growth is likely to be closely related to some weighted sum of LHS variables and so may create simultaneity problems.

\(^9\)We prefer ULCs for the manufacturing sector because they correspond more closely to goods exports than total economy ULCs which also include services and the government sector.

\(^{10}\)The weights are calculated as the mean shares over the 12 previous quarters to avoid endogeneity problems.
To capture the effect of changes in the sectoral composition of trade, we construct an index of sectoral shifts.\(^{11}\)

\[
\Delta C_{i,t}^* = \sum_{k=1}^{K} \phi_{ikt} \Delta S_{ikt}^* \tag{2.4}
\]

where \(\phi_{ikt}\) is the weight of sector \(k\) at time \(t\) for country \(i\), which is determined by a three-year moving average of the share of sector \(k\) in total exports. To avoid simultaneity issues, shares of each good in "world trade" are calculated over all other countries (i.e. excluding the country concerned). Thus \(S_{ikt}^*\) measures the importance of industry \(k\) in the total exports of all other countries in the sample of advanced economies, i.e.:

\[
S_{ikt}^* = \frac{\sum_{j=1, j\neq i}^{I} X_{jkt}}{\sum_{j=1, j\neq i}^{I} \sum_{k=1}^{K} X_{jkt}} \tag{2.5}
\]

Constructed this way, the level variable \(C^*\) measures the general evolution in the relative importance of sectors in the export flows of advanced economies, weighted by the country-specific export shares of the particular sectors. A rise in \(C^*\) indicates that sectoral demand shifts create an increase in the demand for a country’s exports. We opt for a group of 34 OECD countries \((j)\) as the benchmark group for our sample of reporting countries \((i)\), as we want to analyse the compositional effects of world demand that affect advanced economies' exports.\(^{12}\)

In practice, the sectoral shift variable does exhibit significant variation across countries, and in a way which appears orthogonal to trade-weighted GDP growth. By way of

\(^{11}\)An example may help here. Suppose there are two goods in the world, gin and tonic. Initially these are combined in the ratio 1:4 to form the composite drink "G&T". Suppose that the total number of servings of G&T (i.e. world output) is unchanged, but that due to a preference shock consumers now prefer the drink to be mixed with a ratio of 1:3. The share of gin in world trade thus rises by 5 percentage points (from 20% to 25%), the share of tonic in world trade falls by the same amount (from 80% to 75%). Globally, the demand for gin would be 25% higher than before. This benefits countries specialised in the export of gin, at the expense of those specialising in tonic. For a country which only exported gin, the change in demand arising from this shock will be 25%.

\(^{12}\)The 16 reporting countries are a subsample of the reference group of 34 OECD countries.
illustration, the left panel of figure 2.1 shows trade-weighted GDP growth for a selection of countries. Australia and Japan who are more exposed to China and other emerging Asian economies have experienced stronger external demand growth than Europe or the US. But the right hand panel shows that changes in the sectoral composition of trade have affected these two countries very differently. Australia, where commodities, fuel and minerals account for over 60% of exports (more than double any other country in our sample) has benefitted from a shift in trade composition towards these items. Japan, where exports have been geared towards the machinery sectors, has by contrast seen a modest decline in $C^*$ over the sample period. Further details of the coverage and data definitions are given in the appendix.

![Figure 2.1 External growth and sectoral shocks: 2000-2011](image)

3.3 Econometric model specification

The econometric analysis is based on a single equation error correction model (ECM) specification, in line with e.g. Ca’Zorzi and Schnatz (2007), di Mauro et al (2008) and Breuer and Klose (2013). Our variables are expressed in natural logarithms and are within-country demeaned to bring them to the same scale and focus on the time variation within countries. Denoting these transformed variables with lower case letters, our baseline specification is:

$$\Delta x_{i,t} = \alpha(x_{i,t-1} - [\gamma y_{i,t-1} + \beta e_{i,t-1} + \delta c_{i,t-1}]) + \sum_{j=1}^{L_d} \gamma_j \Delta x_{i,t-j}$$

$$+ \sum_{j=0}^{L_e} (\gamma_j^* y_{i,t-j} + \beta_j^* e_{i,t-j} + \delta_j^* e_{i,t-j}) + con_i + \nu_{i,t}$$

(2.6)
The ECM framework permits us to separate out the influence of short-run versus long-run factors on trade. Since volumes may adjust only slowly to changes in relative prices and demand, there are good grounds to believe that the longer-run reaction of exports to a given shock may differ from what happens in the same quarter. In addition, the property that exports tend to revert back to a long-run equilibrium level following a shock may be important in capturing the bounce-back effects of exports witnessed since the financial crisis. The choice of an ECM is furthermore appropriate for the non-stationary level variables given the finding of significant error correcting properties in the model from diagnostic tests (see appendix B for details). The first (round) brackets in equation (2.6) represent the long-run relationship where the long-run coefficients are indicated with the superscript \( l \). The second term captures the short-run dynamics. The average speed of adjustment to the long-run equilibrium is governed by \( \alpha \), the error correction coefficient. To distinguish the short-run coefficients, we depicted them with the superscript \( s \).

Two econometric characteristics of macro panel data sets need to be considered for the choice of the appropriate panel estimator - cross-sectional dependence of the error terms and slope heterogeneity across panel units, as they can both lead to biased coefficient estimates. Cross-sectional dependence in the errors in general reflects factors that are common across countries but are not explicitly accounted for in the model. Not accounting for cross section dependence results in inappropriate standard errors and even in biased coefficient estimates if the common factor in the residuals is correlated with the regressors (Pesaran, 2006). Imposing slope homogeneity on the other hand, can also result in biased coefficients in dynamic panel models when slope heterogeneity is in fact present (Pesaran and Smith, 1995).

To ensure consistent and unbiased estimates, we employ different panel estimators and perform diagnostic tests on the residuals to discriminate between the estimators. We consider the one-way FE estimator next to the Mean Group (MG) and pooled Mean Group (PMG) estimators. The FE estimator assumes all slopes to be homogenous, whereas the latter two estimators do not impose homogeneity on the short-run coefficients. The MG estimator in addition also treats the long-run coefficients to be heterogeneous. The Akaike Information Criterion (AIC) is employed to select the lag length in equation (2.6). The maximum number of lags is restricted to 3 based on the rule \( 4^* (T/100)^{2/9} \) suggested by Gengenbach, Urbain and Westerlund (2008) with \( T=53 \), i.e. the minimum time dimension of the individual country series when no additional lags are included, to preserve a sufficient
number of degrees of freedom. The AIC suggests a lag order choice of one for the FE estimator and three for the MG and PMG estimators, so $L^d = L^e = 1$ and $L^d = L^e = 3$.

The estimation results are shown in table 2.1 for both lag lengths. The Bewley (1979) transformation is used to obtain the long-run coefficients and their standard errors. This is a two-step procedure in which the dependent variable ($\Delta x_{i,t}$) is first regressed on its lagged level, the contemporaneous levels of the exogenous regressors and the differenced terms. In a second step, $x_{i,t}$ is regressed on the first stage fitted value, the contemporaneous levels of the exogenous regressors and the differenced terms. The estimates on the level variables of this second regression provide the long-run coefficients. In case of the MG estimator, the estimated individual long-run coefficients are averaged to obtain the long-run MG coefficients. One can infer from the table that especially the error correction coefficient and the corresponding long-run slope estimates show noticeable differences between the considered panel estimators. Examining the validity of their underlying assumptions is thus important.

We first apply the cross section dependence test (CD test) of Pesaran (2004) to the residuals of equation (2.6) to analyse the extent of correlation between the cross section errors. The CD tests indicate the presence of a substantial amount of cross-sectional correlation in the residuals in both cases (see bottom lines of table 2.1). This outcome is of course not surprising given our discussion in section 2 on the role of globalisation and increased international competition for individual countries’ exports. The inclusion of country-weighted measures of global external demand ($y^*$) and changes in the sectoral composition of advanced economies’ exports ($c^*$) is thus insufficient to account for all common behaviour of exports.

We therefore opt to use Common Correlated Effects (CCE) estimators as these estimators allow us to take account of the influence of unobserved common factors on the coefficient estimates by augmenting the model with the cross section averages (CSAs) of the variables. More specifically, the residual term is specified as follows:

---

13The empirical analysis was carried out in Stata 13, and we employed the user-written Stata routines xted and xtmg written by Markus Eberhardt (Eberhardt, 2012) and xtpmg writtend by Edward Blackburne and Mark Frank (Blackburne and Frank, 2007). The MG results are based on outlier-robust means (by employing the robust option of the xtmg command). To enhance the readability of the table, the estimates for the lagged first difference terms are not shown. This also holds for table 2.2.
\[
\Delta x_{it} = \alpha(x_{i,t-1} - \gamma_t y_{it-1} + \beta_t \epsilon_{it-1} + \delta_t c_{it-1}^s) + \sum_{j=1}^{L_d} \gamma_j \Delta x_{i,t-j} + \sum_{j=1}^{L_e} \gamma_j \Delta y_{i,t-j} + \sum_{j=1}^{L_{c}} \gamma_j \Delta c_{i,t-j}^s + \epsilon_{it} + \nu_{it}
\]

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</tr>
<tr>
<td>Real exchange rate</td>
<td>(e)</td>
<td>-0.123***</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(0.028)</td>
</tr>
<tr>
<td>Sectoral composition</td>
<td>(c^*)</td>
<td>-0.096</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(0.272)</td>
</tr>
<tr>
<td>Error correction</td>
<td></td>
<td>-0.047***</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(0.010)</td>
</tr>
</tbody>
</table>

<table>
<thead>
<tr>
<th></th>
<th>(L^d = L^e = 1)</th>
<th>(L^d = L^e = 3)</th>
</tr>
</thead>
<tbody>
<tr>
<td>World output</td>
<td>(y^*)</td>
<td>2.020***</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(0.005)</td>
</tr>
<tr>
<td>Real exchange rate</td>
<td>(e)</td>
<td>-0.865***</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(0.008)</td>
</tr>
<tr>
<td>Sectoral composition</td>
<td>(c^*)</td>
<td>0.545***</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(0.021)</td>
</tr>
<tr>
<td>CD statistic:</td>
<td>12.09***</td>
<td>9.43***</td>
</tr>
<tr>
<td>average correlation</td>
<td>0.134</td>
<td>0.105</td>
</tr>
</tbody>
</table>

Number of observations: Total=1260, N=16, min T=52, max T=96, average T=79 for a lag order choice of 1.
and Total=1228, N=16, min T=50, max T=94, average T=77 for a lag order choice of 3.

Note: *, **, *** denote significance at 10, 5, and 1% levels respectively. Standard errors are in brackets, p-values in italics.

Table 2.1 Estimation results - standard estimators
where $f_t$ captures an unspecified number of unobserved common factors. Together with the country-specific factor loadings, $\lambda_i$, the term $\lambda_i f_t$ allows to control for cross section dependence and time-variant heterogeneity. The CCE approach augments the model with the CSAs of both the dependent and independent variables in order to account for these unobserved factors. In the particular case of our export model, common factors can be linked to the globalisation of the world economy, increased international competition, general trends in the importance of country blocks in the world economy and common cost shocks to the model. In order to compare the results of the different CCE estimators, table 2.2 shows the results for both lag orders $L^d$ and $L^e$ equal to one and three, based on the AIC.

Table 2.2 shows that the error terms are substantially less correlated across countries when the CCE estimators are employed. The CD-test statistics on the residuals after the inclusion of the CSAs nevertheless remain significant. The statistics and average pairwise correlations are however considerably smaller, which leads us to conclude that the possible bias of the coefficient estimates due to cross-sectional correlated residuals is considerably reduced in light of weak cross section correlation.

A second consideration concerns the possible heterogeneity of the coefficients across countries. The pooled estimator (CCEP) is the most efficient estimator of the CCE estimators but will be inconsistent if the slopes actually differ between the panel units. The same holds for the CCE pooled mean group (CCEPMG) estimator if the long-run slope coefficients actually differ between the panel country units. A look on the coefficient estimates in table 2.2 again suggests that the estimates of the error correction term and the long-run coefficients vary substantially between the estimators. Formal Wald tests can be applied to examine the homogeneity restrictions imposed by the pooled and (pooled) mean group estimators. Table 2.3 displays the F-test statistics and the corresponding p-values for the homogeneity restrictions underlying the different estimators.

Wald tests on the homogeneity assumption of the CCEP versus the CCEMG estimator, suggest a rejection of the restrictions at the 1 per cent level. Similarly, the homogeneity assumption on the short–run slope coefficients of the CCEP estimator relative to the CCEPMG estimator is also rejected. The same holds for the test on the long-run homo-
\[\Delta x = \alpha(x_{t-1} - [\gamma y_{t-1}^j + \beta e_{t-1} + \delta c_{t-1}^j] + \sum_{j=1}^{L} \gamma_j \Delta x_{t-j} + \sum_{j=1}^{L} \delta_j \Delta y_{t-j}^j + \delta_j^c \Delta c_{t-j}^j] + \text{con} + v_t\]

<table>
<thead>
<tr>
<th></th>
<th>(L^d = L^s = 1)</th>
<th>(L^d = L^s = 3)</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>CCEP</td>
<td>CCEPMG</td>
</tr>
<tr>
<td><strong>Short-run coefficients</strong></td>
<td></td>
<td></td>
</tr>
<tr>
<td>World output</td>
<td>(y^*)</td>
<td>0.838** 0.326</td>
</tr>
<tr>
<td></td>
<td>(0.317) 0.008</td>
<td>(0.473) 0.491</td>
</tr>
<tr>
<td>Real exchange rate</td>
<td>(e)</td>
<td>-0.173*** -0.150***</td>
</tr>
<tr>
<td></td>
<td>(0.030) 0.000</td>
<td>(0.038) 0.000</td>
</tr>
<tr>
<td>Sectoral composition</td>
<td>(c^*)</td>
<td>-0.315 -0.298</td>
</tr>
<tr>
<td></td>
<td>(0.281) 0.263</td>
<td>(0.279) 0.286</td>
</tr>
<tr>
<td>Error correction</td>
<td></td>
<td>-0.300 -0.349</td>
</tr>
<tr>
<td></td>
<td>(0.023) 0.051</td>
<td>(0.060) 0.030</td>
</tr>
<tr>
<td><strong>Long-run coefficients</strong></td>
<td></td>
<td></td>
</tr>
<tr>
<td>World output</td>
<td>(y^*)</td>
<td>0.974*** 0.700***</td>
</tr>
<tr>
<td></td>
<td>(0.071) 0.000</td>
<td>(0.189) 0.000</td>
</tr>
<tr>
<td>Real exchange rate</td>
<td>(e)</td>
<td>-0.767*** -0.637***</td>
</tr>
<tr>
<td></td>
<td>(0.019) 0.000</td>
<td>(0.034) 0.000</td>
</tr>
<tr>
<td>Sectoral composition</td>
<td>(c^*)</td>
<td>-0.124** 0.011</td>
</tr>
<tr>
<td></td>
<td>(0.063) 0.047</td>
<td>(0.122) 0.038</td>
</tr>
<tr>
<td>CD statistic:</td>
<td></td>
<td>-5.55*** -4.48***</td>
</tr>
<tr>
<td>average correlation:</td>
<td></td>
<td>-0.060 -0.048</td>
</tr>
</tbody>
</table>

Number of observations: Total=1248, N=16, minT=52, maxT=94, average T=78 for a lag order of 1; Total=1228, N=16, minT=50, maxT=94, average T=77 for a lag order of 3. Note: **, *** denote significance at 10, 5, and 1% levels respectively. Standard errors are in brackets, p-values in italics. Note that the p-values for the error correction term are not listed, given that the t-statistic is not standard normally distributed due to the approximation of unobserved common factor by the CSAs. The significance levels for the CCEMG estimator can instead be based on the simulated critical values provided by Gengenbach et al (2008, see appendix B).

Table 2.2 Estimation results - CCE estimators
Tests of homogeneity restrictions of the CCEPMG estimator versus the CCEMG estimator. Based on these tests, the CCEMG estimator can be considered to be the preferred estimator.

Phillips and Sul (2003) however show that Wald tests are not reliable in a dynamic panel setup with cross section dependence. Given the finding of a low extent of cross-section dependence under the CCE estimators, the results should be interpreted with care. A Hausman test on the estimated coefficients of the CCEMG and CCEPMG estimators however also reveals that the obtained long-run panel coefficients are significantly different. The test statistic is 18.65 with a p-value of 0.000 for a lag order of one and 10.49 with a p-value of 0.015 for a lag order of three and suggests that the consistent estimator, CCEMG, is to be preferred over the efficient estimator, CCEPMG. The finding of cross-country variation in the slope coefficients of the export demand model is furthermore in line with the evidence for cross-country variation in the literature (e.g. Carlin et al, 2001; di Mauro et al, 2008).

Given the results of the cross-section dependence and homogeneity tests, we use the CCEMG estimator for the remainder of our empirical analysis. The CCEMG estimator is shown to be consistent in a dynamic single equation model (Chudik and Pesaran, 2013) under the assumptions that the regressors are weakly exogenous, that the time dimension of the panel is sufficiently long and that the number of CSAs corresponds to the number of unobserved factors. To save degrees of freedom and to avoid over-parametrization, a general-to-specific method was employed on the CCEMG estimates based on the method proposed by Hendry (1993). Starting from 3 lags in the dynamics, any insignificant lags were stepwise excluded. This procedure resulted in a parsimonious specification for the CCEMG estimator which includes the contemporaneous first differences of the exogenous regressors, except for the sectoral variable for which up to two lags of the first differences

<table>
<thead>
<tr>
<th></th>
<th>CCEP versus CCEPMG</th>
<th>CCEP versus CCEMG</th>
<th>CCEPMG versus CCEMG</th>
</tr>
</thead>
<tbody>
<tr>
<td>statistic</td>
<td>1.54</td>
<td>2.29</td>
<td>3.60</td>
</tr>
<tr>
<td>p-value</td>
<td>0.000</td>
<td>0.000</td>
<td>0.000</td>
</tr>
</tbody>
</table>

\( L^d = L^c = 1 \)

<table>
<thead>
<tr>
<th></th>
<th>CCEP versus CCEPMG</th>
<th>CCEP versus CCEMG</th>
<th>CCEPMG versus CCEMG</th>
</tr>
</thead>
<tbody>
<tr>
<td>statistic</td>
<td>1.30</td>
<td>2.58</td>
<td>2.46</td>
</tr>
<tr>
<td>p-value</td>
<td>0.005</td>
<td>0.000</td>
<td>0.000</td>
</tr>
</tbody>
</table>

\( L^d = L^c = 3 \)

Table 2.3 Tests of homogeneity restrictions
remain included in the model. This results in the following empirical model specification:

\[
\Delta x_{i,t} = \alpha_i(x_{i,t-1} - [\gamma_i y_{i,t-1} + \beta_i l_{i,t-1} + \delta_i l_{i,t-1} + \lambda_i l_{i,t-1}]) + \\
\sum_{j=0}^{2} (\gamma_i \Delta y_{i,t} + \beta_i \Delta l_{i,t} + \delta_i \Delta l_{i,t} + \lambda_i \Delta l_{i,t}) + con_i + \varepsilon_t
\]  

(2.8)

where \( l_{i,t} \) and \( l_{i,t}^s \) denote the CSAs for the variables in the long-run and short-run part of ECM.

4 Estimation results

The estimation results of specification (2.8) are shown in table 2.4. Column I shows the baseline specification, with output, real exchange rate and the sectoral composition variable appearing in both the long-run and short-run parts of the ECM. Regression II explores the possibility of differential reactions to the two components of the real exchange rate. Separating these out reveals a clear difference in the short run—the elasticity with respect to the nominal exchange rate is 0.05, but the elasticity with respect to unit labour costs is about six times as large. A formal Wald test rejects (p-value of 0.06) the null that the two coefficients are equal. This finding supports the theoretical reasoning that one could expect a larger impact of relative unit labor costs to export prices due to the comovement of nominal exchange rates and imported inputs and due to their larger persistence relative to the nominal exchange rate. The point estimates of the two coefficients also show a notable difference in the long run and the corresponding Wald test rejects their equality as well, although the p-value is close to the 10% boundary (p-value of 0.09). The unequal effects of the REER components thus continue to hold in the long run.

The coefficients have the expected signs. An expansion in world GDP by 1% leads on average to a 0.89% expansion in a country’s exports, while the long-run response is 1.9%. The short-run elasticity of exports with respect to the nominal exchange rate is not significant but has a significant long-run coefficient of -0.21. Similarly, the short-run elasticity of exports with respect to the relative ULCs is 0.30, rising to 0.42 in the long run. The values of these coefficients are in line with earlier work.\(^{14}\) The contemporaneous

\(^{14}\)E.g. Bayoumi, Harmsen and Turunen (2011) and Chen \textit{et al} (2013)
### Chapter 2

#### Table 2.4 Estimation results - benchmark model (2.8)

<table>
<thead>
<tr>
<th>Estimation sample:</th>
<th>I 1984Q1-2008Q</th>
<th>II 1984Q1-2008Q</th>
<th>III 1984Q1-2012Q1</th>
</tr>
</thead>
<tbody>
<tr>
<td><strong>Short-run coefficients</strong></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>World output $y^*$</td>
<td>0.519 (0.407)</td>
<td>0.886* (0.495)</td>
<td>0.852 (0.520)</td>
</tr>
<tr>
<td>Real exchange rate $e$</td>
<td>-0.098** (0.045)</td>
<td>-</td>
<td>-</td>
</tr>
<tr>
<td>Nominal exchange rate $s$</td>
<td>-</td>
<td>-0.051 (0.056)</td>
<td>-0.089* (0.046)</td>
</tr>
<tr>
<td>Unit labour costs $u$</td>
<td>-</td>
<td>-0.300** (0.117)</td>
<td>-0.288** (0.121)</td>
</tr>
<tr>
<td>Sectoral composition $c^*$</td>
<td>-0.008 (0.416)</td>
<td>0.419 (0.563)</td>
<td>0.779* (0.409)</td>
</tr>
<tr>
<td>$c_{t-1}$</td>
<td>0.669 (0.408)</td>
<td>0.475 (0.459)</td>
<td>0.170** (0.408)</td>
</tr>
<tr>
<td>$c_{t-2}$</td>
<td>0.791*** (0.197)</td>
<td>0.665** (0.270)</td>
<td>0.512** (0.229)</td>
</tr>
<tr>
<td>Error correction</td>
<td>-0.571 (0.050)</td>
<td>-0.643 (0.056)</td>
<td>-0.584 (0.067)</td>
</tr>
<tr>
<td><strong>Long-run coefficients</strong></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>World output $y^*$</td>
<td>1.695*** (0.404)</td>
<td>1.947*** (0.480)</td>
<td>1.892** (0.521)</td>
</tr>
<tr>
<td>Real exchange rate $e$</td>
<td>-0.366*** (0.088)</td>
<td>-</td>
<td>-</td>
</tr>
<tr>
<td>Nominal exchange rate $s$</td>
<td>-</td>
<td>-0.209** (0.088)</td>
<td>-0.331** (0.065)</td>
</tr>
<tr>
<td>Unit labour costs $u$</td>
<td>-</td>
<td>-0.415*** (0.081)</td>
<td>-0.546** (0.082)</td>
</tr>
<tr>
<td>Sectoral composition $c^*$</td>
<td>0.180 (0.318)</td>
<td>0.144 (0.346)</td>
<td>0.317 (0.360)</td>
</tr>
<tr>
<td>CD statistic:</td>
<td>-4.41*** (0.571)</td>
<td>-4.33*** (0.677)</td>
<td>-4.30*** (0.578)</td>
</tr>
<tr>
<td>average correlation:</td>
<td>-0.047</td>
<td>-0.046</td>
<td>-0.042</td>
</tr>
</tbody>
</table>

Number of observations for respectively columns I and II and column III: Total=1276, N=16, min T=53, max T=97, average T=80 and Total=1516, N=16, min T=68, max T=112, average T=95. Note: *, **, *** denote significance at 10, 5, and 1% levels respectively. Standard errors are in brackets, p-values in italics. Note that the p-values for the error correction term are not listed, given that the t-statistic is not standard normally distributed due to the approximation of unobserved common factor by the CSAs (see appendix).
sectoral composition variable is not significant although the lagged first differences have a significant positive effect on exports. The error correction term is strongly significant, with a value of -0.64, implying that any disequilibrium is corrected over less than two quarters. The results for the sample including the GTC (1984Q1-2012Q1) are further displayed in column III. Comparing the second and third column, it is clear that the estimates for the pre-GTC and entire period are fairly similar.

Given the results of the Wald tests on the REER components, our preferred specification allows for different effects of the components both in the short run and the long run. The model estimates in column II are by consequence used to construct the dynamic forecasts in the next section.

5 Conditional forecasts

We use the model to conduct a dynamic forecast exercise in order to better understand export performance since the GTC. Based on the CCEMG estimates of our preferred model, we examine how the forecasts compare to actual export outturns to infer whether the countries’ exports are in line with the model predictions if one considers the expected average effect of the traditional demand and competitiveness determinants. The sample under analysis is the pre-GTC period (1984Q1 to 2008Q2), with the forecast period beginning at the turning point in global trade (2008Q3).

First, we focus on a country-average export forecast to analyse whether the model’s predictions are in line with actual observations during and following the GTC. In the first instance, we construct an out-of-sample forecast conditional on the actual paths of the external demand, competitiveness and sectoral variables. Afterwards, the forecasts are conditional on the realised values of the macroeconomic determinants of exports as well as on the observed values of the CSAs. This forecast thus includes both the exogenous shocks to external demand, sectoral composition and real exchange rates, as well as the additional information about common unobserved shocks which is embodied in the CSAs.

Next, we examine the individual forecasts to evaluate the performance of each country relative to its peers. The country-specific forecasts represent a benchmark for what country exports "should have been" based on the obtained CCEMG coefficient estimates and given

\[15\] As with any dynamic forecast, we forecast the change in each quarter and then compute a levels forecast by cumulating these forecasts in changes over time.
the evolution of the country-specific exogenous variables and the CSAs. The forecast for example allows to capture the general performance of other countries because each country’s forecast utilises information about the cross section averages of all countries. By using average coefficient estimates, we obtain a uniform benchmark to compare each country’s relative export performance. Our measure of performance is defined as the gap between actual and forecast exports, expressed as a percentage of forecast exports. A positive value indicates that a country has outperformed the forecast based on the average coefficients, a negative value indicates underperformance.

5.1 Aggregate trade

To understand what the forecasts imply for the path of the advanced economy exports as a whole, we combined the individual country data and forecasts into weighted averages based on each country’s average share of exports between 2008Q1 and 2011Q4. Figure 2.2 shows the profiles of these variables. The dashed line represents actual exports and the solid lines the out-of-sample forecasts. The dotted lines depict the in-sample performance of the model. The solid line with the squared markers shows the aggregate forecast when the influence of the CSAs is not taken into account. To remove the influence of the CSA terms during the forecast period, the CSAs of the level variables are kept at their value of 2008Q2 and the CSAs of the differences are set to zero since 2008Q3. The second solid line depicts the aggregate forecast that in addition takes into account the observed values of the CSAs.

From figure 2.2, we can infer that the model approximates average exports since the start of the GTC quite well. At the time of the GTC itself, actual exports fell faster and farther than what the model would have predicted out of sample when ignoring the influence of the CSAs and explains only about half of the fall in exports between the peak in 2008Q2 and the through in 2009Q2. The out-of-sample forecast conditional on the CSAs however almost fully captures that fall from peak to through. This difference is an indication that global factors, represented by the combined effect of all the included variables their CSAs and not captured by the model’s country-specific export determinants, were a key driver of the GTC. Since the through, both out-of-sample forecasts evolve closer to each other. The out-of-sample forecasts on average show a level difference

16These shares were calculated on the basis of the nominal US dollar value of exports for each country and reflect the share of country i in the total exports of all 16 countries.
of about 1 and 0.5 \% to actual exports since 2010Q1, respectively when conditioning on both the country-specific variables and the CSAs and when not including the CSAs. The in-sample forecast based on the CCEMG coefficients (red dotted lines) overpredicts the fall in exports between 2008Q2 and 2009Q2 by 18\% and reaches an average difference of about 3.5\% to actual exports since 2010Q1 as it starts to diverge from actual exports from 2011 onwards. When the individual-specific coefficients (CCEMG\(_i\), blue dotted line) are used, the in-sample forecast explains 97\% of the fall from peak to through and approaches actual export dynamics after the GTC to a larger extent. This suggests that the largest part of the deviation from the in-sample forecast based on the average coefficients to actual exports can be explained by the use of the average instead of individual-specific coefficients.

![Figure 2.2 Actual weighted average of exports versus forecast](image)

Figure 2.2 Actual weighted average of exports versus forecast

To gain more insight in the model’s dynamics, we computed a decomposition of the profile of the out-of-sample forecast that conditions on the CSAs, shown in figure 2.3. The decomposition clearly shows that the combined effect of the CSAs is the driving factor that brings the forecast down. Since the through in actual exports, the effects of changes in country-specific external demand (shown by the yellow bars) start to push down the forecast while the effects of the CSAs become less dominant. There is a small boost from competitiveness after 2009Q4 (the grey bars) but sectoral shifts continue to have small
downward effect (pattern bars).

Figure 2.3 Forecast decomposition (weighted average)

Taken together, this suggests that the explained fall in exports during the GTC can be mainly attributed to the evolution of the common factors outside the scope of models at the individual country level. Given that the aggregate out-of-sample forecast conditioning on the observed CSAs captures the evolution of average actual exports since the GTC to a large extent, the individual country forecasts based on the CCEMG coefficients can offer a relevant benchmark to evaluate individual export performances since the GTC.

5.2 Evaluating individual countries performance

The forecast profiles for each country conditional on the observed country-specific variables and the CSAs, together with the actual outturn are shown in figure 5.2. This figure clearly indicates that the level of actual exports relative to the forecasted exports varies substantially across the countries in our panel. Based on the CCEMG coefficients, some countries (e.g. Germany and Netherlands) show better export outcomes than their benchmark whereas other countries’ exports are below the benchmark (e.g. Japan and Canada). For other countries (e.g. Portugal and Italy) actual exports seem to be well approximated by the forecast. So, given the CSAs and the country-specific evolutions in the
traditional demand and competitiveness determinants, the advanced economy countries’ export performance since the GTC relative to their benchmark differs to a considerable degree.

Figure 2.4 Actual exports versus forecasts for individual countries

The measures of relative performance are shown below in figures 2.5 and 2.6, for respectively 2009Q2 and 2011Q4. We compare actual versus forecast exports at these two points in time to evaluate the export performance in the immediate aftermath of the GTC in 2009Q2 and next in 2011Q4, to gauge the longer-run effects of the GTC on trade dynamics. As already discussed above, the level of exports is on average close to the model’s prediction in both periods.

This average however reveals considerable variation across countries. On this metric, the strongest performers are Austria, the Netherlands and Germany with exports well above 10% more than the model would have predicted in both quarters. Actual exports of the European countries are in general better relative to the forecast benchmark, although France also underperforms relative to its forecast in both quarters. Actual exports of individual countries thus relate to a different extent to what can be expected based on
Figure 2.5 Actual exports versus benchmark, 2009Q2

Figure 2.6 Actual exports versus benchmark, 2011Q4
the panel’s average coefficients of the traditional external demand and competitiveness variables. Conditional on the average coefficients estimates, the United Kingdom (UK) has underperformed by around 10% in 2009Q2 and by 8% in 2011Q4. Whilst the UK’s cumulative growth in exports from 2008Q1 to 2011Q4 is close to the cross-country median, the large depreciation in sterling (and accompanying fall in the ULC based real exchange rate) failed to boost exports by as much as the model would have predicted.

6 Conclusions

We analyse what has happened to advanced economies’ exports since the global Great Trade Collapse witnessed during 2009Q2 and 2008Q3. To this end, we estimate a dynamic panel model of goods exports for 16 advanced economies that includes the traditional export determinants, being foreign demand and competitiveness. We incorporate a measure of the sectoral composition of trade as an additional determinant of the demand for exports and we allow for a different effect of nominal exchange rates and relative costs on exports. In addition, by using the Common Correlated Effects estimator of Pesaran (2006) we control for unobserved common factors such as common globalisation dynamics and evolutions in the degree of international competition. We find that the two components of the real effective exchange rate, the nominal exchange rate and relative unit labour costs, have a significant different effect on exports. The sectoral composition variable has only a small influence on export dynamics which is confined to a lagged effect in the short run.

We use the model to construct a forecast benchmark to evaluate countries’ export performance based on the average panel estimates of the model prior to the Great Trade Collapse. First, we consider how the pre-crisis model fared in predicting the overall path of advanced economy exports. We find that the model can only explain around half of the fall in exports during the Great Trade Collapse based on country-specific evolutions in the variables but subsequently approximates actual exports. If the common unobserved factors, proxied by the cross section averages of all variables, are taken into account, the evolutions of exports are well fitted over the entire forecast period. This finding suggests that common factors played a major role during the Great Trade Collapse.

Given this good average forecast performance, we evaluate individual countries’ export performance against their peers by constructing a forecast conditioning on both the path of the country-specific variables and common unobserved factors. We find substantial
variation across countries in terms of their actual exports relative to the forecasts based on the average panel coefficients. For the United Kingdom in particular, exports in 2011Q4 came in about 8 per cent below the benchmark suggested by international comparisons.
Appendices

A. Data sources and construction of variables

Real Exports: chain volume index of exports in national currencies, seasonally adjusted (OECD Quarterly National accounts). UK export data are corrected for MTIC fraud.

Trading partners’ GDP: real GDP volume, seasonally adjusted, (Datastream, ??GDP...D code, except for Spain [code ESGDP..VE], Brazil [code BRGDP...G], Poland [code POSMGDPC], Estonia [code EOGDPPIPDF], China [IMF World Economic Outlook data], Malta [Eurostat data], the CensusX12 procedure is applied if the series are not yet seasonally adjusted)

-> sample of countries for construction trade-weighted global GDP: 76 countries

[ UK, Austria, Belgium, Estonia, Finland, France, Germany, Greece, Iceland, Ireland, Italy, Luxembourg, Netherlands, Portugal, Slovak Republic, Slovenia, Spain, Cyprus, Malta, United States, Japan, Canada, Australia, New Zealand, Denmark, Norway, Sweden, Switzerland, Singapore, South Korea, Albania, Bulgaria, Croatia, Czech Republic, Hungary, Latvia, Lithuania, Poland, Romania, Serbia, Turkey, Jordan, Belarus, Georgia, Moldova, Russia, Ukraine, Argentina, Brazil, Chile, Colombia, Costa Rica, Ecuador, Mexico, Peru, South Africa, China, India, Indonesia, Malaysia, Philippines, Thailand, Israel, Kenya, Namibia, Uganda, Venezuela, Egypt, Iran, Morocco, Tunisia, Paraguay, Mozambique, Sri Lanka ]

Bilateral export flows, for different sectors: goods exports for 63 sectors, current US dollar (OECD International Trade by Commodity Statistics database, Standard International Trade Classification (SITC) Revision 2)

-> sample of countries for advanced economies group: 34 OECD countries

[ Australia, Austria, Belgium, Canada, Chile, Czech republic, Denmark, Estonia, Finland, France, Germany, Greece, Hungary, Iceland, Ireland, Israel, Italy, Japan, South Korea, Luxembourg, Mexico, Netherlands, New Zealand, Norway, Poland, Portugal, Slovak Republic, Slovenia, Spain, Sweden, Switzerland, Turkey, United Kingdom, United States ]
### Table 2.5 Overview number of observations per country

- 63 sectors: 69 available sectors - sector 35, 91, 93, 95, 96 and 97 due to data quality (large amount of missing observations)
- Annual data: 1980-2012, interpolated to quarterly data

**Effective exchange rates:** nominal and real ULC-based effective exchange rate indices, double-weighted to capture import and export competition in third markets (IMF, International Financial Statistics). For more details on the methodology, see Bayoumi, Lee and Jayanthi (2005).
-Chapter 2-

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Table 2.6 Average cross-sectional correlation of residuals ADF tests

B. Time series properties

We apply the cross section dependence test of Pesaran (2004) to the residuals of individual augmented Dicky-Fuller (ADF) tests.\(^\text{17}\) The results signify substantial cross section residual correlation between the country-specific series (see table 2.6), which calls for the use of second generation panel unit root tests which take into account cross-sectional dependence and provide critical values at the panel level which are simulated such that they are valid under dependence of the individual series.

Given the unbalanced nature of our dataset, we opt for the cross-sectionally augmented IPS (CIPSM) test of Pesaran, Smith and Yamagata (2013).\(^\text{18}\) The extent of cross-sectional dependence of the residuals from the individual cross section augmented ADF tests (CADF) based on this approach, is substantially reduced for all variables (table 2.7, bottom part). The CIPSM tests further indicate that the variables $x$, $y^*$, $e$ and $c^*$ can considered to be nonstationary, although the test results on the $y^*$ and $e$ variables depend on the imposed lag order.

To find out whether equation (2.8) constitutes a meaningful long-run relation, we apply the panel error correction test of Gengenbach, Urbain and Westerlund (2008). This approach tests the significance of the error correction term within a conditional ECM framework that allows for possible nonstationary common factors. If the error correction term is found to be significant, this implies the existence of a long-run equilibrium relationship. Two panel tests of the null hypothesis of no error correction that the error

\(^\text{17}\) A constant and trend are included for $x$, $y^*$, a constant only for $c^*$ and $e$, both in the ADF and CIPSM tests

\(^\text{18}\) The cross-sectional averages of the other variables in (2.1) are added as additional common factors for the CIPSM tests on $x$, $y^*$, $e$ and $c^*$. 

66
Table 2.7 Average cross-sectional correlation of residuals CIPS M tests

correction term for every $i$ are provided, given by the (truncated) average of individual t-tests and Wald tests. Based on the results of the unit root tests, we test the following reparametrized conditional ECM-specification:

$$\Delta x_{it} = \alpha_i (x_{i,t-1} - [\beta_{1,i} y_{i,t-1} + \beta_{2,i} e_{i,t-1} + \beta_{3,i} c_{i,t-1}]) + \sum_{j=1}^{L} \beta_{4,ij} \Delta x_{i,t-j} + \sum_{j=0}^{L} (\beta_{5,ij} \Delta y_{i,t-j} + \beta_{6,ij} \Delta e_{i,t-j} + \beta_{7,ij} \Delta c_{i,t-j}) + \sum_{j=0}^{L} \psi_{ij} \tilde{CA}_{t-j} + \rho_{it} \quad (2.9)$$

where $\tilde{CA}$ stands for the cross section averages which are used as proxies for unobserved common factors. The t-test directly tests whether $\alpha_i$ is significantly different from zero whereas the Wald tests test whether $\alpha_i$ and the coefficients on the lagged levels of the exogenous regressors are jointly equal to zero. The panel tests’ null hypotheses are both rejected at the 5% significance level for a lag order choice of 4, respectively with a statistic of -3.88 and 34.96 where the critical values at the 1% significance level and for $N=20$ are respectively-3.53 and 20.36. Also for a lag order choice of 3, both tests reject the null with statistics of respectively -4.12 and 34.62.\(^{19}\) The individual t-tests and Wald tests cannot

\(^{19}\)The lag order $L$ is determined based on the Aikaike Information Criterion, where the maximum number of lags is set equal to $4*(T/100)^{2/9}$ as suggested in Gengenbach et al (2008). This resulted in a maximum of 4 lags and a lag order choice of 4 for average and maximum $T$ and in a maximum of 3 lags and a lag order choice of 3 for the minimum $T$. 

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average residual cross section dependence

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Note: *, **, *** denote significance at 10, 5, and 1% levels respectively.
reject the null for respectively 9 and 6 of the 16 countries for both lag order choices. The panel test thus displays evidence of error correction, whereas the individual tests cannot reject the null for an important subset of countries. Taking into account that the power of the tests are greatly improved by pooling (Gengenbach et al, 2008), we conclude that including the level information in our variables next to their differences is appropriate.
Bibliography


Chapter 2


CHAPTER 3

Inflation during times of economic slack and deleveraging: a panel data analysis.

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Abstract

Historically, persistent and pronounced economic downturns gave rise to notable falls in the level of inflation but these falls are not observed in the aftermath of the global financial crisis of 2008-2009. This paper analyzes inflation dynamics in a cross-country Phillips curve framework while considering credit evolutions and periods of financial stress and documents a flattening of the Phillips curve during economic slack.

JEL classification: E31, E32, E51

Keywords: Phillips curve, bank credit, financial crisis, cross-country panel

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1 Introduction

Most advanced economies are today recovering from a pronounced economic downturn following the global financial crisis of 2008-2009. These downturns were particularly large in historical terms but were also special in the sense that they occurred in association with a synchronized financial crisis. In addition, the immediate aftermath of the financial crisis has been marked to contain a "missing disinflation puzzle" since advanced economies’ inflation rates after the financial crisis have declined less than what one would expect based on standard Phillips curve coefficients given the large and often very persistent activity gaps.\(^1\) Commonly heard explanations to rationalize the missing disinflation puzzle within the standard framework are incorrect output gap measures, a better anchoring of inflation expectations, the low level of inflation and increased globalization. An alternative avenue is to refer to unobservable shocks to the markup pricing behavior of firms. But also nonlinearities and asymmetries in the Phillips curve relationship could account for the muted reaction of inflation.

The association of the downturns with a financial crisis however prompts one to question whether financial distortions have influenced the subsequent inflation dynamics beyond their impact via real output. A financial crisis typically involves impaired bank balance sheets and a reduced intermediation role of banks. Consistent with standard credit channels, deflationary pressures are expected to arise following restrained credit conditions. But next to the indirect effect of credit on inflation via GDP, credit rationing could also alter the price-setting behavior of firms and, in turn, inflation directly as firms can be forced to look for alternative and possibly more expensive funding sources. Gilchrist et al (2014) show in a calibration of a general equilibrium model that aggregate inflation dynamics are indeed directly influenced by changes in financial conditions when financial frictions are allowed to enter the model. These authors find that especially under severe financial distortions, firms put off investments in their market share via a lowering of the price in response to changes in their cash flow or financing conditions.\(^2\) The change in the price-setting behavior may then counteract deflationary pressures driven by increased spare capacity, especially during periods of financial stress. Accordingly, the relation be-

\(^1\)Ball and Mazumder (2011) for example show this for the United States.

\(^2\)This model implication is consistent with the empirical findings in Gilchrist et al (2014) on price-setting behavior of large publicly-traded US firms based on income, balance sheet and product-level price data during 2005 and 2012.
between inflation and economic activity might be affected by the association with financial frictions and this can influence the Phillips curve estimates.

This paper relates to the recent literature on the relationship between credit and business cycles, which received a vast amount of attention in light of the Great Recession. Previous findings\(^3\) point out that the typical shape of a business cycle that is characterized by a financial and/or debt crisis differs from the shape of a standard business cycle. The run-up in debt during the expansion phase is typically higher, the recession deeper and longer, while the recovery tends to be more sluggish than for a normal cycle due to weak private demand and credit conditions. Bijsterbosch and Dahlhaus (2011) furthermore show that recoveries without a pick up in credit are more likely after more severe recessions while their incidence even doubles after financial crises. These creditless recoveries moreover occur at a slower pace than recoveries associated with credit growth (Bijsterbosch and Dahlhaus, 2010; Abiad, Dell’Ariccia and Li, 2011) although Bech, Gambacorta and Kharroubi (2012) and Takáts and Upper (2013) find that deleveraging is not always detrimental for the subsequent recovery. This is particularly the case if the run-up to a financial crisis is characterized by a high growth in credit, which can reveal a credit boom. To the extent that the accumulation of debt during a credit boom leads to excessive debt levels or a low quality of debt, deleveraging of the private sector may be neutral or even supportive for the subsequent recovery. One can thus conclude that the source of the recession and the subsequent credit conditions matter for the course of real GDP. It is however not obvious whether credit evolutions also directly influence inflation dynamics and the Phillips curve relationship.

The focus in this paper is therefore twofold: first, we analyze whether the reaction of inflation to economic activity differs depending on the sign, magnitude and persistence of the deviation from potential activity. This analysis is warranted given the substantial evidence of nonlinearities and asymmetries in the Phillips curve depending on the level of economic activity documented in the existing literature\(^4\) and their potential to account for the missing disinflation puzzle.

Second, we examine whether the reaction of inflation to economic activity is in addition affected by the credit cycle. We particularly concentrate on credit cycle downturns and

\(^3\)Terrones, Scott and Kannan (2009), Claessens, Kose and Terrones (2012), Leigh et al (2012)  
\(^4\)E.g. Laxton, Meredith and Rose (1995); Turner (1995); Barnes and Olivei (2003); Stock and Watson (2010).
upturns next to the occurrence of banking crises. As such, we try to answer the question whether the large extent of economic slack can explain the relatively mild disinflation during and in the aftermath of the global financial crisis and whether the association with a financial crisis and related credit evolutions attenuated the inflation reaction.

In addition, we shed light on the possibility of a speed limit effect during periods of spare capacity driven by a bounce-back in output. The occurrence of a high-growth recovery after a deep recession, the so-called bounce-back effect, has been documented by, amongst others, Beaudry and Koop (1993), Sichel (1994) and Piger, Morley and Kim (2005). The bounce-back occurs when there is no new production capacity needed to meet recovering demand in presence of unused capacity. A quick pick up in aggregate demand could however trigger speed limit effects (Turner, 1995; Dwyer, Lam and Gurney, 2010) that give rise to inflationary pressures because resources cannot be reactivated immediately. Speed limit effects could be more pronounced during periods of large and persistent spare capacity due to a larger extent of obsolete and inadequate production capacity. A bounce-back in GDP might thus result in a milder inflation reaction during economic slack episodes. Information on the existence of speed limit effects and their possible nonlinear nature is especially relevant to infer inflation dynamics when aggregate demand recovers, which is the situation currently faced by most advanced economies.

We analyze inflation dynamics in a hybrid New Keynesian Phillips curve framework for a panel of 15 OECD countries. Financial crises are rare events and persistently large negative deviations of output from its potential level do not occur that often either. This presents a challenge for the empirical analysis of the effect of economic activity on inflation during these episodes. A cross-country perspective offers a way to obtain a sufficient number of observations on large negative output gaps, financial crises, credit downturns and their combinations.

We find activity gaps to be significant drivers of inflation and obtain evidence for an asymmetric inflation reaction. Inflation reacts considerably more to economic activity when production capacity falls short of demand relative to when capacity is below potential and in the latter scenario, there is no significant inflation reaction. The estimation results further suggest that there is no additional effect of the credit cycle on the slope of the hybrid New Keynesian Phillips curve. Similarly, we cannot establish that the Phillips curve slope is affected during recessions and recoveries associated with periods of financial distress. Our findings imply that the occurrence of economic slack, as experienced nowadays by
many advanced economies, would result in a relatively limited reaction of inflation to output. Credit evolutions do not influence inflation dynamics over and above their indirect effect on aggregate output. A lower downward pull on inflation during these economic slack periods is thus expected, but credit evolutions do not seem to directly affect the inflation response to economic activity.

The remainder of the paper is organized as follows: section 2 outlines the employed Phillips curve model and briefly reviews the possible drivers of nonlinearities and asymmetries in the Phillips curve. Section 3 describes how we define business and credit cycles and their mutual relations while section 4 explains the empirical strategy and presents the estimation results. Section 5 offers some robustness checks and section 6 concludes.

2 Examining inflation dynamics through the lens of the Phillips curve relationship

2.1 Hybrid New Keynesian Phillips curve

We focus on the estimation of a hybrid New Keynesian Phillips curve (HNKPC) as the Phillips curve allows us to examine the effect of the level of economic activity on inflation next to inflation expectations while controlling for other short-term pressures. The structural formulation of the New Keynesian Phillips curve (NKPC) implies that inflation is purely forward looking as price formation depends on the expected evolution of demand and supply factors. We instead employ a hybrid form of the Phillips curve to control for the observed persistence in realized inflation, in which case the model is augmented with lagged inflation terms that capture inertia in the formation of inflation expectations. We assume a proportional relation between firms’ marginal costs and economic activity, such that the employed specification of the Phillips curve is expressed in terms of economic activity instead of (unobserved) real marginal costs.

A general representation of the HNKPC is given by the following specification:

\[
\pi_{i,t} = \alpha_i + \gamma \text{gap}_{i,t} + \lambda E_t[\pi_{i,t+\tau}] + \sum_{j=1}^{l} \rho_j \pi_{i,t-j} + \sum_{j=0}^{g} \varphi_j \pi_{i,t-j} + \varepsilon_{i,t} \tag{3.1}
\]

Where \(\pi\) represents the inflation measure\(^5\), \(\text{gap}\) the level of spare capacity, \(E[\pi_{t+\tau}]\)

\(^5\)Denoting the price index by \(p_t\), inflation is defined as \(\pi_t = 100\% \left( \frac{p_t}{p_{t-1}} - 1 \right)\).
inflation expectations, and ext external supply shock variables. \( \alpha_i \) denotes the country fixed effects and the subscripts \( i \) and \( t \) denote countries and time periods. The disturbance term \( \varepsilon_{i,t} \) represents measurement errors and any other combination of unobserved shocks.

### 2.2 Possible originators of asymmetries and nonlinearities

**Level of economic activity** To analyze whether the combination of large economic slack and credit evolutions induced by a financial crisis entails an atypical reaction of inflation to economic activity, we first focus on specific segments of the Phillips curve. The reaction of inflation to economic activity has been found to be asymmetric and nonlinear. Laxton et al. (1995) and Turner (1995) document that high levels of economic activity increase inflation more than that low levels of activity reduce it. Barnes and Olivei (2003) and Stock and Watson (2010) on the other hand find the Phillips curve to be essentially flat during normal times with relatively small activity gaps whereas the curve steepens in periods of large gaps. A more muted inflation reaction since 2008-2009 is in line with the first finding of asymmetric behavior whereas the second seems to confirm the missing disinflation hypothesis as economies were confronted with large negative output gaps, in which case you would expect a larger reaction of inflation. Asymmetric nonlinearities however offer a possible explanation that merges both findings in the sense that the nonlinearities could mainly manifest themselves during expansions.

**Credit evolutions** We next consider the evolution of credit to the private sector as a possible explanation for a flattening Phillips curve. The reaction of inflation to economic activity might be altered due to changes in the firms’ price-setting behavior driven by external financing conditions, especially in times of financial distress (Gilchrist et al., 2014). Gilchrist et al. (2014) posit that changes in financial conditions can directly affect inflation dynamics in presence of financial frictions. To incorporate these frictions, the authors depart from the standard assumption of frictionless financial markets and base their model on the theory of customer markets (Bils, 1989). Under the theory of customer markets, firms actively set prices as an investment tool to build their future customer base. As such, a firm has the incentive to lower prices today to increase its market share in the (near) future. However, when external finance is costly or the firms’ cash flow is low, firms might decide to reduce the extent of investment in the customer base via price reductions and increase their price. This mechanism is therefore capable to alter the
reaction of aggregate price dynamics to contractionary demand side shocks during periods of financial distortions. When aggregate financing conditions worsen, the aggregate price level will be pushed upwards which counteracts deflationary pressures.

In the first instance, we focus on the influence of downturns and upturns in the credit cycle on the HNKPC estimates. Downturns in the credit cycle are not necessarily corresponding to periods of severe financial distortions but can nevertheless pick up periods of more stringent funding conditions that influence macroeconomic outcomes. An important caveat in this respect is that country-specific aggregate credit data do not allow us to distinguish between the part of private sector deleveraging that is due to supply constraints and the part that reflects a lower demand for credit triggered by contractions in economic activity. We therefore analyze the effects of the credit cycle on inflation dynamics while controlling for the business cycle and the level of capacity utilization.

As a second test of the influence of restrictive credit conditions on the Phillips curve slope, we elaborate upon the reaction of inflation during recessions and recoveries that occur in correspondence with times of outright financial stress. We employ banking crises datings as a key indicator of severe financial turmoil and restrained credit supply. Banking crisis start dates allow us to investigate the repercussions of a full-blown financial crisis, as during 2008-2009, on the estimation of the Phillips curve relationship in the subsequent quarters. Given that the end dates of the crisis dating are to some extent subjective in nature and are therefore prone to discussion, we also examine credit boom periods as an alternative indicator of pronounced financial stress. Schularick and Taylor (2012) namely underline that the fragility of a financial system is related to credit evolutions. The authors show that the risk of a financial crisis during a certain period is significantly increased when a credit boom occurs over the previous five years. Credit booms preceding a recession and subsequent recovery may therefore also signal the emergence of episodes of elevated financial stress throughout the credit bust. By focusing on the association of the business cycle phases with periods of financial stress, we in addition examine whether the Phillips curve slope is particularly sensitive to financial stress depending on the phase of the business cycle.

In conclusion, inflation could react differently to economic activity depending on the sign, the level and the persistence of the activity gap. The inflation reaction might further be directly affected by financial distortions.
2.3 Speed limit effects

We further quantify the importance of speed limit effects, as these effects might change inflation dynamics for a given level of the output gap. We specifically focus on a pick up in aggregate demand measured by positive changes in the output gap when the gap is negative, i.e. on periods when the output gap "closes". We verify whether the speed limit effects are influenced by the severity and persistence of the negative deviation of output from its potential level.

3 Business and credit cycle definitions and banking crises dating

Before we turn to the estimation results of the HNKPC, this section briefly describes the definitions which were used to construct the different phases of economic activity, the credit cycle and banking crisis episodes. A more elaborate description of the employed definitions is given in appendix C.

**Persistent economic slacks/booms** We consider economic slack periods to be large when the negative output gap is larger than a threshold value of 1.5% and we quantify them to be persistent if this lasts for at least 8 quarters, based on Meier (2010). Economic boom periods are similarly defined based on positive deviations of output from potential. Persistent economic boom periods thus capture periods of at least 8 consecutive quarters with an output gap above 1.5%.

**Business cycle phases** We define business cycle expansions and recessions based on a threshold on the cyclical component of output similar to the method applied in the work of, amongst others, Sugawara and Zalduendo (2013). Expansions run from trough to peak (excluding the trough) and recessions run from peak to trough (excluding the peak). The threshold value that is used to identify business cycle troughs equals minus the standard deviation of the output gap. A peak is next determined based on the largest value of

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6The 1.5% threshold is also in line with the estimated threshold values for the unemployment gap in Barnes and Olivei (2003) for the US. These are respectively -1.34 and 1.40 for the core CPI inflation measure and a time-varying natural rate of unemployment.
Banking crises, bank credit cycle and bank credit booms  We follow the dating of banking crises in Drehmann, Borio, and Tsatsaronis (2011) and Laeven and Valencia (2013) to infer the start dates of banking crises. To date credit cycles, we employ the classical methodology used in Claessens et al (2012) on the real credit per capita data. This classical approach is based on the algorithm put forward by Harding and Pagan (2002) and allows us to quantify upturns and downturns in credit based on absolute declines and increases in real credit per capita. The algorithm considers the value $y_t$ to be a peak if $(y_t - y_{t-2}) > 0$, $(y_t - y_{t-1}) > 0$, $(y_{t+2} - y_{t}) < 0$ and $(y_{t+1} - y_t) < 0$, while a trough occurs at time $t$ if $(y_t - y_{t-2}) < 0$, $(y_t - y_{t-1}) < 0$, $(y_{t+2} - y_{t}) > 0$ and $(y_{t+1} - y_t) > 0$.

We also apply thresholds on the bank credit data to identify credit boom periods. Credit booms are defined as consecutive dates for which the cyclical component of the Hodrick-Prescott filtered log of real credit per capita is equal or larger than 1.5 times its standard deviation. The peak of a credit boom is subsequently defined as the quarter from the set of consecutive quarters that satisfy the credit boom condition with the largest difference between the cyclical component and the threshold standard deviation. Given the peak, the start and end date of the boom are the quarters before and after the peak with the minimum difference between the cyclical component and its standard deviation.

It should be noted that we limit our attention to bank credit data, obtained from the BIS bank credit database. A credit boom driven by non-bank financial intermediaries would consequently not be captured in the data that we use. Limited data availability and comparability hamper a cross-country analysis of alternative funding sources and with the exception of the United States (US), bank credit accounts for the dominant share of total credit to the private sector. The BIS database is nevertheless particularly attractive as it offers a measure of total bank credit provided by both domestic and foreign banks, which is an important advantage compared to data on credit from domestic banks only given the increased importance of foreign banks in domestic markets.

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8 We use the BIS series on total credit to the private non-financial sector at quarterly frequency (see appendix for details).
Restrictive credit conditions under different business cycle phases Bank crises are assumed to be associated with recessions if the crisis occurs during the 8 quarters before the start of the recession or during the recession itself.\(^9\) In the same vein, we consider recession episodes that are preceded by a credit boom during the 8 quarters preceding the recession to indicate periods of elevated financial stress. Economic recoveries are deemed to be associated with a banking crisis or to be preceded by a credit boom if these episodes occur within the 12 quarters preceding the expansion. We further define recoveries to be creditless following Sugawara and Zalduendo (2013) if the average growth rate of real bank credit per capita during the recovery is below or equal to zero.

Results identification methods Using these methodologies on the underlying data series since 1960Q1, we identify 45 recessions periods and 40 expansions in our sample of 15 countries since 1990 of which respectively 28 and 33 are complete business cycle phases. We further identify 41 credit downturns and 44 credit upturns (of which 33 and 27 are completed), 14 credit booms and 18 creditless recoveries. The precise identification of the business and credit cycle phases inevitably demands some judgement. We therefore apply an alternative business cycle definition in the robustness section in addition to already considering different credit evolution indicators.

By way of illustration, figures 3.1 and 3.2 outline the results of the identification methodologies for the US and Italy. The bars depict the values of the output gap and make the distinction between recessions and expansions. The shaded background areas denote credit cycle downturns. Periods of persistent and large economic slack are also highlighted (dotted bars) next to creditless recoveries (bold horizontal striped bars). Both figures illustrate that the number of persistent and large economic slack periods and the number of severe recessions associated with a banking crisis or preceding credit boom is rather small for individual countries.

Table 3.1 displays the details on the precise dates of these large and persistent slack periods for the entire sample of countries, the interaction with the defined business and credit cycles, and their association with a periods of restrictive credit conditions. Over the 15 countries, we identify 24 periods of large and persistent output gaps where only South Korea did not experience such episode. Columns 3 and 4 indicate that 20 of the

\(^9\)This definition is based on Bech, Gambacorta and Kharroubi (2012).
Figure 3.1 Evolution output gap US and defined business and credit cycle episodes

Onset banking crisis: 1990Q2 and 2007Q3  
credit boom: 2007Q2-2009Q2

- recession  - expansion  - persistent economic slack  - creditless recovery

Figure 3.2 Evolution output gap Italy and defined business and credit cycle episodes

Onset banking crisis: 1992Q3 and 2008  
credit booms: 1991Q3-1992Q4, 2007Q3-2008Q4

- recession  - expansion  - persistent economic slack  - creditless recovery
### Table 3.1 Persistent economic slack periods

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<td>yes</td>
<td>yes</td>
<td>yes</td>
<td>no</td>
</tr>
<tr>
<td>Finland 1991q2-1997q1 / 2009q1-2010q4</td>
<td>yes / no / yes / no</td>
<td>yes / no</td>
<td>yes / no</td>
<td>yes / yes</td>
</tr>
<tr>
<td>France 2008q4-2010q4</td>
<td>yes / yes</td>
<td>yes</td>
<td>yes</td>
<td>yes / yes</td>
</tr>
<tr>
<td>Germany 2003q1-2005q4 / 2009q1-2010q4</td>
<td>yes / yes / yes / yes</td>
<td>yes / no</td>
<td>no / no</td>
<td>yes / yes</td>
</tr>
<tr>
<td>Ireland 1992q3-1995q1 / 2008q4-2013q1</td>
<td>no / yes / no / no</td>
<td>no / yes</td>
<td>no / no</td>
<td>no / no</td>
</tr>
<tr>
<td>Italy 1992q4-1994q4 / 2009q1-2013q1</td>
<td>yes / yes / yes / yes</td>
<td>yes / yes / yes / yes</td>
<td>no / yes</td>
<td>yes / yes</td>
</tr>
<tr>
<td>Korea, South -</td>
<td>-</td>
<td>-</td>
<td>-</td>
<td>-</td>
</tr>
<tr>
<td>Netherlands 2003q1-2005q1</td>
<td>no / no</td>
<td>no / no</td>
<td>yes / no</td>
<td>no</td>
</tr>
<tr>
<td>Norway 1993q2 / 2009q1-2011q2</td>
<td>yes / yes / yes / yes / yes</td>
<td>yes / no</td>
<td>no / no</td>
<td>no / no</td>
</tr>
<tr>
<td>Sweden 1992q1-1995q3 / 2008q4-2010q4</td>
<td>yes / no / yes / no</td>
<td>yes / yes</td>
<td>yes / no</td>
<td>yes / no</td>
</tr>
<tr>
<td>US 2008q4-2013q1</td>
<td>yes / yes</td>
<td>yes</td>
<td>yes</td>
<td>yes</td>
</tr>
</tbody>
</table>

Total number of:
- Credit downturns: 20
- Banking crises: 17
- Completed credit booms: 7
- Creditless recoveries: 14
24 persistent economic slack periods are associated with a credit cycle downturn\textsuperscript{10}, either during the recession or expansion part of these persistent slack periods or during both. The next columns offer some more information on the association of the business cycles phases with periods of restrictive credit conditions during the persistent economic slack periods.

17 of the persistent economic slack periods are associated with a banking crisis, again during the recession or expansion part of the persistent slack periods or during both. 7 of the persistent economic slack periods are preceded by a completed credit boom and 14 recoveries during these persistent economic slack periods are defined to be creditless. From this table, one can conclude that a large amount of persistent economic slack periods are also identified as periods of elevated financial stress. This finding highlights the importance of analyzing whether inflation dynamics are affected by financial stress over and above the influence through aggregate output.

4 Empirical analysis

4.1 Linear Phillips curve representation

We consider the following linear specification of the HNKPC in the remainder of this work:

\[
\pi_{i,t} = \alpha_i + \gamma gap_{i,t} + \sum_{j=1}^{k} \delta_j \Delta gap_{i,t-j} + \lambda E_t[\pi_{i,t+4}] + \sum_{j=1}^{l} \rho_j \pi_{i,t-j} + \sum_{j=0}^{g} \varphi_j \widehat{\pi}_{i,t-j} + \varepsilon_{i,t} \quad (3.2)
\]

Where \(E_t[\pi_{t+4}]\) denotes one-year-ahead inflation expectations and \(\Delta gap\) the first difference of the \(gap\) variable. \(\Delta gap\) enters equation (3.2) to measure speed limit effects, i.e. the effects of lagged changes in economic activity on inflation for a given level of economic activity.

We opt for core inflation as the inflation measure to capture general price developments while abstracting from temporary fluctuations. The model can of course be easily extended to headline inflation. Economic activity is measured by the OECD output gap

\textsuperscript{10}Note that we consider a credit downturn to be associated with a recession and expansion similar to the criterion employed for banking crises and credit booms; i.e. if the downturn occurs in the 8 quarters before or during the recession and if the downturn occurs during the 12 quarters preceding the expansion.
series. The measure is equal to the percentage difference between the levels of actual GDP and estimated potential GDP and allows to control for time-variation in potential output. Admittedly, the assessment of the level of spare capacity in real time is challenging and subject to important ex-post revisions but this challenge is considered to be a distinct research topic. In this work, the downward pull on inflation of spare capacity levels is analyzed conditional on the available (OECD) output gap data.

An alternative activity measure is based on the work of Beaudry and Koop (1993) who define a depth of recession variable as the difference between the current level of output and its historical maximum.\textsuperscript{11} This measure is developed in order to take possible nonlinearities depending on the magnitude of deviations from output from its time-varying maximum into account. We likewise construct an output recession gap measure calculated as the difference of real GDP and its maximum over the current and 11 previous quarters, expressed in percentages of that maximum. This measure by consequence only considers the timing of economic contractions and their severity.

We use the survey-based Consensus Economics one-year-ahead inflation forecasts data as a direct measure of inflation expectations following, amongst others, Roberts (1997). Empirical evidence shows that survey measures are in general not rational, which violates the rational expectation assumption on the formation of expectations underlying the standard New Keynesian Phillips Curve framework. Henzel and Wollmershäuser (2008) however derive the HNKPC under the assumption of subjective expectations and find the specification to be identical to the specification of the HNKPC specification derived under rational expectations. In addition, the authors demonstrate that the use of survey data gives more reliable empirical results than the instrumental-variable-based rational expectations approach. The drawback of abandoning the rational expectation assumption is that the HNKPC specification is no longer microfounded (Mavroeidis, Plagborg-Møller and Stock, 2014) because the formation of expectations is left unmodelled. Following the standard approach in empirical work on the Phillips curve that uses survey expectations, we treat the survey inflation forecasts as exogenous for the estimation of the HNKPC. We thereby implicitly assume that \( \varepsilon_{t,t} \) is a pure news shock such that the use of contemporaneous (and therefore not predetermined) forecasts is valid.\textsuperscript{12}

\textsuperscript{11}Stock and Watson (2010) similarly define an unemployment recession gap measure as the difference between the current unemployment rate and its minimum over the current and 11 previous quarters.
\textsuperscript{12}Alternatively, one might consider lagged survey expectations, i.e. \( E_{t-1}[\pi_{t,t+4}] \), since these are definitely predetermined. The estimates in table 3.2 are however highly similar in this case (the results are available
We further include two external price shock variables ($ext$) as additional cost factors for a given level of economic activity. We consider in this respect the dynamics of import prices of commodity goods ($ext^c$) and of import prices of non-commodity goods and services ($ext^{nc}$). Both series are normalized such that a zero value can be interpreted as the absence of an external price shock. The data sources for the series are listed in the appendix.

We allow for a general downward trend in the level of inflation by estimating equation (3.2) alternatively in inflation gap form, where $\pi_{i,t}$, its lags and the one-year-ahead expectations are replaced by their deviation from long-term inflation expectations, i.e. $(\pi_{i,t-j} - E_t[\pi_i^T])$ and $(E_t[\pi_{i,t+4}] - E_t[\pi_i^T])$. The long-term inflation expectations are thus used to measure the trend level of inflation. This specification allows for a drift in the average inflation rate over time as in Faust and Wright (2012) and previous works. We use the Consensus Economics six-to-ten-year-ahead inflation forecasts data for the long-term inflation expectations where available and extend them with the long-run Hodrick-Prescott filtered trend of headline inflation since 1960Q1 or the first available quarter of the series.

We estimate this HNKPC relationship for an unbalanced panel of 15 OECD countries over the period 1990Q1-2013Q1. We thereby posit that the number of time observations is sufficiently large to avoid the Nickell (1981) bias in a dynamic panel data setup. We limit our attention to OECD countries to have a group of countries which are relatively similar in institutional and structural characteristics. The group of countries is further determined by the data availability at quarterly frequency. The sample period further starts in 1990 due to the cross-country data availability of inflation expectations but is also suited since we want to exclude the end of the Great Inflation period, which runs from the beginning of the 1970s until the mid-1980s, and focus on episodes of considerable financial market globalization and integration.

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from the author on request). Because it is notoriously hard to find strong instruments that satisfy the exclusion restriction, an instrumental variable approach is not considered in this work.

13 e.g. Kozicki and Tinsley (2001), Cogley, Primiceri and Sargent (2010), Stock and Watson (2010) and Clark (2011). The specification assumes the unpredictability of changes in trend inflation.

14 These are: Australia, Canada, Denmark, Finland, France, Germany, Ireland, Italy, Japan, Korea, Netherlands, Norway, Sweden, United Kingdom and the United States. The extent of unbalancedness is very limited as the sample only starts later for Germany, Ireland and South Korea (respectively 1991Q3, 1991Q1 and 1991Q2) due to the data availability while the end dates are balanced.
4.2 Estimation results

When considering macro panel estimators, one needs to check the validity of the underlying assumptions of the estimators. In particular, the possible heterogeneity of the slope coefficients and the presence of cross section dependence in the error terms driven by common unobserved variables needs to be tested. Neglecting heterogeneity of the slopes will result in biased coefficient estimates in a dynamic panel (Pesaran and Smith, 1995) while the presence of unobserved common factors can lead to inconsistent variance estimates and even to inconsistent coefficients if the factors are correlated with the variables in the model (Pesaran, 2006).

We first examine the slope homogeneity assumption underlying the standard Fixed Effects (FE) estimator relative to the Mean Group (MG) estimator that allows for heterogeneous coefficients by testing the restrictions by means of the Swamy test. Table 3.2 shows the estimation outcomes for equation (3.2) based on the different panel estimators and the outcome of the corresponding Swamy test on the slope homogeneity assumption of the FE estimator (see bottom lines).\(^{15}\) The statistic is highly significant and thus suggests the rejection of slope homogeneity.

Second, we test for cross-sectional correlation in the error terms by means of the cross section dependence test (CD-test) of Pesaran (2004). To this end, we employ the Common Correlated Effects Mean Group (CCEMG) estimator developed by Pesaran (2006) which allows to control for common unobserved factors in the residuals by including the cross section averages of all variables to the model. The CCEMG estimator is shown to be consistent in a dynamic single equation model (Chudik and Pesaran, 2013) under the assumptions that the regressors are weakly exogenous, the time dimension of the panel is sufficiently long and the number of unobserved factors is adequately captured by the cross section averages. The CD-test statistics are displayed below in table 3.2 and indicate the presence of significant correlation in the residuals of the model estimated by the FE and MG estimators but an insignificant correlation for the CCEMG estimator.

---

\(^{15}\)The estimated equation includes the contemporaneous term of the output gap and one-year-ahead inflation expectations, 2 lags of the first difference of the output gap, 5 lags of inflation and the contemporaneous term and 3 lags of the commodity goods and non-commodity goods and services import price inflation variables. So \(k=2, l=5\) and \(g=3\) in (3.2). This lag structure is based on the significance of the variables when \(k, l\) and \(g\) increase from 1 to 8 (where the maximum lag is determined by the Akaike information criterion) to obtain a parsimonious model.
Given the results of the homogeneity test and the CD-tests, the coefficients in the third column of table 3.2 are considered as benchmarks values for the estimates of the nonlinear Phillips curve specifications in section 4.4. Allowing for heterogeneous slopes and proxies for unobserved common factors especially affects the estimates of the backward-looking expectation component.

The slope is significant and implies that there exists a positive empirical link between the output gap and inflation measures, although the magnitude is relatively modest. A change in the output gap of 1% leads to an increase in quarter-on-quarter inflation of 0.03%. The speed limit effect, measured by the sum of the $\delta$ coefficients, is not significantly different from zero according to an F-test. The estimates of $\lambda$ and $\sum_{j=1}^{5} \rho_j$ are positive and significantly different from zero, suggesting the combination of forward-looking and backward-looking components of inflation dynamics. The sum of the $\lambda$ and $\sum_{j=1}^{5} \rho_j$ estimates turns out to be significantly different from 1 and equals 0.64. This finding indicates that the dynamic process is stable, i.e. the sum does not exceed one. The fourth column shows the CCEMG results for the inflation gap form. Controlling for a slowly-varying trend inflation, however, does not change the estimation results of (3.2) in an important way.

4.3 Incorporating nonlinearities and asymmetries

To assess whether the reaction of inflation to economic activity is indeed different depending on the level of economic activity and credit conditions and whether speed limit effects are relevant, we extend equation (3.2) to the following multiplicative interaction model:

$$
\pi_{i,t} = \tilde{\alpha}_i + \gamma gap_{i,t} + \sum_{j=1}^{k} \delta_j \Delta gap_{i,t-j} + \lambda E_t[p_{i,t+4}] + \sum_{j=1}^{l} \rho_j \pi_{i,t-j} + \sum_{j=0}^{g} \varphi_j \tilde{c}xt_{i,t-j} + \varepsilon_{i,t} \quad (3.3)
$$

where

$$
\tilde{\gamma} = \gamma_1 I(D_{i,t}^{gap}) + \gamma_0 \\
\tilde{\delta} = \delta_1 I(D_{i,t}^{\Delta gap}) + \delta_0 \\
\tilde{\alpha} = \alpha_1 I(D_{i,t}^{con}) + \alpha_0 \\
\tilde{D}_{i,t}^{con} = (D_{i,t}^{gap}, D_{i,t}^{\Delta gap})
$$

\footnote{For the FE and MG estimators, the sum is not significantly different from 1, also suggesting a stable process.}
\[ \pi_{t,t} = \alpha_i + \gamma_{i} \text{gap}_{t,t} + \sum_{j=1}^{k} \delta_j \Delta \text{gap}_{t,t-j} + \lambda E_t[\pi_{t,t+4}] + \sum_{j=1}^{l} \rho_j \pi_{t,t-j} + \sum_{j=0}^{g} \gamma_j \text{t}_{t,t-j} + \xi_{t,t} \]

### Table 3.2 Estimation results - linear Phillips curve

<table>
<thead>
<tr>
<th></th>
<th>( \gamma )</th>
<th>( \delta )</th>
<th>( \lambda )</th>
<th>( \rho_j )</th>
<th>( \rho_{t,j} )</th>
<th>( \rho_{t,j}^{nc} )</th>
<th>CD-test</th>
<th>Average correlation</th>
</tr>
</thead>
<tbody>
<tr>
<td>(1) FE</td>
<td>0.018***</td>
<td>0.026*</td>
<td>0.535***</td>
<td>0.471***</td>
<td>0.021</td>
<td>0.0221</td>
<td>2.32**</td>
<td>0.024</td>
</tr>
<tr>
<td>(2) MG</td>
<td>0.018***</td>
<td>-0.009</td>
<td>0.620***</td>
<td>0.331***</td>
<td>0.008</td>
<td>0.0088</td>
<td>3.25***</td>
<td>0.033</td>
</tr>
<tr>
<td>(3) CCEMG</td>
<td>0.030**</td>
<td>0.011</td>
<td>0.469***</td>
<td>0.171**</td>
<td>0.016</td>
<td>0.0168</td>
<td>1.60</td>
<td>0.016</td>
</tr>
<tr>
<td>(4) CCEMG</td>
<td>0.021</td>
<td>0.091</td>
<td>0.434**</td>
<td>0.144*</td>
<td>-0.016</td>
<td>0.034</td>
<td>1.53</td>
<td></td>
</tr>
</tbody>
</table>

CD-test statistic: 2.32**  3.25***  1.60  1.53

Swamy’s heterogeneity test on FE coefficients

Note: \(*, **, ***\) denote significance at 10, 5, and 1% levels respectively. Standard errors are in brackets, p-values in italics. The hypothesis that the sum of coefficients is equal to zero is tested by means of an F-test. The total number of observations is 1379, with \(T=92\), max \(T=93\), min \(T=86\) and \(N=15\).
Equation (3.3) allows for changes in the level and the slope of the Phillips curve depending on the indicator variables $I(D^\text{con}_{i,t})$, $I(D^\text{gap}_{i,t})$ and $I(D^\Delta^\text{gap}_{i,t})$. $I(D^\text{gap}_{i,t})$ and $I(D^\Delta^\text{gap}_{i,t})$ equal 1 if the specific gap and $\Delta$ gap characteristic one is testing for occurs during the quarter $t$ and equals 0 otherwise. The intercept, the coefficient on the economic activity gap and the coefficient on the change in the gap are thus allowed to be different according to the values of $I(D^\text{gap}_{i,t})$ and/or $I(D^\Delta^\text{gap}_{i,t})$. Our main interest lies in the estimates of $\gamma$ and $\sum \bar{\delta}$ in equation (3.3), which gives the values of the slope of the Phillips curve and the speed limit effect for different values of $D^\text{gap}_{i,t}$ and $D^\Delta^\text{gap}_{i,t}$.\footnote{The standard errors of the estimates of $\alpha$, $\gamma$ and $\bar{\delta}$ are computed via the Delta method. In case of multiple indicator variables, the $\alpha_1$ and $\gamma_1$ or $\delta_1$ coefficients and $D^\text{gap}_{i,t}$ or $D^\Delta^\text{gap}_{i,t}$ are replaced by vectors (which are denoted by a tilde).}

In the next section, we incorporate these dummy variables into the model to analyze whether the $\gamma$ and $\delta$ coefficients in equation (3.2) are affected by the different states of the economy.

### 4.4 Estimation results

#### Level and persistence of output gaps

Table 3.3 presents the results of the analysis of asymmetries and nonlinearities in the Phillips curve depending on the sign, level and persistence of the output gap. Column 1 repeats the benchmark results of estimating equation (3.2) whereas column 2 provides a first test of asymmetric behavior by adding the interactions with the dummy $D^\text{gap}_{i,t}$ equal to one if the output gap is nonnegative and zero otherwise. The difference between $\gamma_0$ and $\bar{\gamma}$, indicated by $\gamma_1^{\text{sign}}$, is significant and implies that the slope coefficient is smaller for negative than for nonnegative output gaps. For negative output gaps, the Phillips curve slope is not significantly different from zero.

Column 3 shows the effects of large output gaps relative to moderate gaps. We allow for an asymmetric threshold effect by including two binary dummies to quantify the effects of output gaps below -1.5 and above 1.5 next to the sign dummy. $\bar{\gamma}_1$ thus captures three slope effects, $\bar{\gamma}_1 = (\gamma_1^{\text{sign}}, \gamma_1^{LP}, \gamma_1^{LN})$, where $\gamma_1^{\text{sign}}$ denotes the change in the slope for positive output gaps, $\gamma_1^{LP}$ for large positive output gaps and $\gamma_1^{LN}$ for large negative gaps. In this case, the interactions are not significant.
\[ \pi_{i,t} = \bar{\alpha}_i + \tau g_{0}p_{i,t} + \sum_{j=1}^{T} \beta_j \Delta g_{0}p_{i,t-j} + \lambda E_{i}[\pi_{i,t+4}] + \sum_{j=1}^{T} \rho_j \pi_{i,t-j} + \sum_{j=1}^{T} \varphi_{j} \varepsilon_{i,t-j} + \varepsilon_{i,t} \]

<table>
<thead>
<tr>
<th>Linear model</th>
<th>Interaction model</th>
<th>Linear model</th>
</tr>
</thead>
<tbody>
<tr>
<td>(1)</td>
<td></td>
<td>(2)</td>
</tr>
<tr>
<td>( D_{g_{0}}^{gap} = 1 ) if: ( g_{0}p_{i,t} \geq 0 )</td>
<td>( g_{0}p_{i,t} &lt; -1.5 ) / ( g_{0}p_{i,t} &gt; 1.5 )</td>
<td>recession gap</td>
</tr>
<tr>
<td>( \gamma )</td>
<td>( \gamma_0 )</td>
<td>( \gamma_0 )</td>
</tr>
<tr>
<td>( \gamma )</td>
<td>( \gamma_1 )</td>
<td>( \gamma_1 )</td>
</tr>
<tr>
<td>( \gamma_{\text{sign}} )</td>
<td>( \gamma_{1}^{\text{sign}} )</td>
<td>( \gamma_{1}^{\text{sign}} )</td>
</tr>
<tr>
<td>( \gamma_{L_{1}N} / \gamma_{L_{1}P} )</td>
<td>( \gamma_{1}^{L_{1}N} / \gamma_{1}^{L_{1}P} )</td>
<td>( \gamma_{1}^{L_{1}N} / \gamma_{1}^{L_{1}P} )</td>
</tr>
<tr>
<td>( \gamma_{\text{slack}} / \gamma_{\text{boom}} )</td>
<td>( \gamma_{1}^{\text{slack}} / \gamma_{1}^{\text{boom}} )</td>
<td>( \gamma_{1}^{\text{slack}} / \gamma_{1}^{\text{boom}} )</td>
</tr>
<tr>
<td>( \sum_{j=1}^{T} \delta )</td>
<td>( \sum_{j=1}^{T} \delta )</td>
<td>( \sum_{j=1}^{T} \delta )</td>
</tr>
</tbody>
</table>

Note: * indicates significance at 10%, ** at 5%, and *** at 1% levels respectively. Standard errors are in brackets, p-values in italics. The hypothesis that the sum of coefficients is equal to zero is tested by means of an F-test. The total number of observations is 1379, with \( T=92 \), max \( T=93 \), min \( T=86 \) and \( N=15 \) but reduces to 1358, with \( T=91 \), max \( T=93 \), min \( T=75 \) when using the recession gap variable.

Table 3.3 Estimation results - level and persistence of output gap (CCEMG)

Column 4 further explores the possibility of differential effects during persistently large output gaps on inflation during either positive or negative output gap values, so \( \gamma_1 = (\gamma_{1}^{\text{sign}}, \gamma_{1}^{\text{boom}}, \gamma_{1}^{\text{slack}}) \). The slope for persistent economic slacks is significantly larger (\( \hat{\gamma}_1 = 0.07 \)) than for the moderate or non-persistently large slack periods (\( \hat{\gamma}_0 = -0.02 \)). The \( \gamma_0 \) estimate points to an insignificant Phillips curve slope for the latter periods. The total slope effect during persistent slack periods nevertheless remains insignificantly different from zero. Persistent economic booms do not generate a significantly different slope estimate. The next column displays the results of the linear model for the output recession gap variable.\(^{18}\) The reaction of inflation to economic activity during periods with output below the time-varying maximum is not found to be significant. The recession gap measure therefore also suggests that a downward pull on inflation does not occur in times of spare capacity which confirms the finding in column 2. This alternative activity gap measure however has the drawback that it imposes that economic activity only exerts an influence on inflation in times of spare capacity. We therefore opt to employ the dummy

\(^{18}\)Note that the time dimension of the sample is reduced (total observations is 1358) because the recession gap variable requires data up to 11 lags of the output gap.
interactions for the level of the output gap instead of imposing such restriction.

In conclusion, these results suggest an asymmetric Phillips curve which is flat for negative output gaps and positively sloped for periods of activity above potential. Persistent economic slack periods do not imply a meaningfully different Phillips curve slope than periods with small to moderate gaps. Speed limit effects are not found to be relevant. Given the evidence of asymmetries in the Phillips curve and the lack of a threshold level effect, we further control for the sign of the output gap in the remainder of this work and do not continue to include the threshold dummies.

**Credit evolutions**

In this section, we focus on the sensitivity of the Phillips curve slope coefficient to bank credit evolutions next to the sign of the output gap. Given that the countries included in the analysis experienced negative gaps throughout 2009Q1-2013Q1 with the exception of Germany and Japan which experienced respectively 7 and 1 quarter(s) with positive values as well\(^{19}\), a focus on the Phillips curve slope estimates for negative output gap values is most appropriate to analyze the missing disinflation hypothesis. We therefore only show the \( \hat{\gamma} \) estimation outcomes obtained for the negative output gap values. First, we analyze the effects of the general credit cycle\(^{20}\) and afterwards we particularly concentrate on restrictive credit conditions depending on the phase of the business cycle. Table 3.4 displays the results for the former.

The first column shows the benchmark estimates of the model in equation (3.3) when only the dummy interaction for the sign of the output gap is included. The next two columns focus on the results of estimating equation (3.3) where the interactions capture upturns in the credit cycle next to business cycle interactions\(^{21}\), so \( \hat{\gamma}_1 = (\gamma_1^{sign}, \gamma_1^{bus}, \gamma_1^{cred}) \) where \( \gamma_1^{bus} = 1 \) during expansions and \( \gamma_1^{cred} = 1 \) during upturns of the credit cycle. Since we only focus on the slope estimates for the negative output gap values, the displayed result for \( \hat{\gamma} \) do not consider the effects of \( \gamma_1^{sign} \). The results indicate that both credit upturns and credit booms do not exercise a significant influence on the Phillips curve.

\(^{19}\)i.e. during 2011Q1-2011Q4 and after 2012Q2 for Germany and in 2013Q1 for Japan.

\(^{20}\)Note that we cannot test for a different slope during banking crises as the datings pinpoint the starting dates of a banking crises but remain vague about the end dates.

\(^{21}\)The business cycle phases are introduced in the estimation of equation (3.3) to control for correlations between business and credit cycles (infra, p. 79).
\[
\pi_{i,t} = \bar{\alpha}_t + \bar{\gamma} \text{gap}_{i,t} + \sum_{j=1}^{k_j} \delta_j \Delta \text{gap}_{i,t-j} + \lambda E_t[\pi_{i,t+4}] + \sum_{j=1}^{\rho_j} \pi_{i,t-\gamma_j} + \sum_{j=0}^{\theta_j} \tilde{\varphi}_{j,t} \text{ext}_{i,t-j} + \varepsilon_{i,t}
\]

**Interaction model**

<table>
<thead>
<tr>
<th>(1)</th>
<th>(2)</th>
<th>(3)</th>
</tr>
</thead>
<tbody>
<tr>
<td>(D_{i,t}^{gap} = 1^a ) if: ( \text{gap}_{i,t} \geq 0 / \text{expansion} )</td>
<td>( \text{gap}_{i,t} \geq 0 / \text{expansion/credit upturns} )</td>
<td>( \text{gap}_{i,t} \geq 0 / \text{expansion/credit booms} )</td>
</tr>
<tr>
<td>( \hat{\gamma}_0 )</td>
<td>0.016</td>
<td>0.026</td>
</tr>
<tr>
<td></td>
<td>(0.017)</td>
<td>(0.020)</td>
</tr>
<tr>
<td>( \hat{\gamma}_1 ) for ( \text{gap} &lt; 0 )</td>
<td>-</td>
<td>0.007</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(0.032)</td>
</tr>
<tr>
<td>( \hat{\gamma}_{bus} )</td>
<td>-</td>
<td>-0.021**</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(0.010)</td>
</tr>
<tr>
<td>( \hat{\gamma}_{cred} )</td>
<td>-</td>
<td>0.002</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(0.022)</td>
</tr>
<tr>
<td>( \sum_{j=1}^{\theta_j} \delta_j )</td>
<td>0.006</td>
<td>0.003</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(0.830)</td>
</tr>
</tbody>
</table>

Note: \(*, **, ***\) denote significance at 10, 5, and 1% levels respectively. Standard errors are in brackets, p-values in italics. The hypothesis that the sum of coefficients is equal to zero is tested by means of an F-test. The total number of observations is 1379, with \( T = 92 \), \( \max T = 93 \), \( \min T = 86 \) and \( N = 15 \).

\( a \) \( \bar{\gamma} \) depicts the slope estimate when the output gap is negative

**Table 3.4 Estimation results - business cycle and credit evolutions (CCEMG)**

slope once controlled for the business cycle. A split up of the business cycle in recessions and expansions leads to a significant interaction term in column 2 but the slope estimate remains insignificant during recessions as well.

We next focus on the association of recessions and recoveries (i.e. the first 8 quarters of expansions) with periods of elevated financial stress, captured by the association of recessions and recoveries with banking crises or preceding credit booms. For the recoveries, we additionally examine the HNKPC slope during creditless recoveries.

**Recessions** In table 3.5, we zoom in on recession periods in combination with periods of severe financial distortions. In this case, \( \tilde{\gamma}_1 = (\gamma_{1}^{\text{sign}}, \gamma_{1}^{\text{bus}}, \gamma_{1}^{\text{fin}}) \) where \( \gamma_{1}^{\text{fin}} = 1 \) during periods of financial stress associated with recessions. Note that \( \tilde{\gamma} \) in table 3.5 now refers to the HNKPC slope estimates when \( \tilde{\gamma}_1 = (0, 0, 1) \) as we are interested in the slope effects during recession periods conditional on credit conditions. Column 2 focuses on the association with a banking crisis and column 3 on credit boom busts. The estimates indicate that the Phillips curve slope coefficient is not significantly affected when recessions are associated with a banking crisis or preceded by a credit boom.
\[
\pi_{i,t} = \bar{\alpha}_i + \tau_{gap_{i,t}} + \sum_{j=1}^{k} \delta_j \Delta gap_{i,t-j} + \lambda E_t[\pi_{i,t+4}] + \sum_{j=1}^{r} \rho_j \pi_{i,t-j} + \sum_{j=0}^{q} \gamma_j E_t t_{i,t-j} + \epsilon_{i,t}
\]

<table>
<thead>
<tr>
<th>Interaction model</th>
<th>(gap_{i,t} \geq 0)</th>
<th>(gap_{i,t} &gt; 0)</th>
</tr>
</thead>
<tbody>
<tr>
<td>(D_{i,t}^{gap} = 1^{a}) if: (gap_{i,t} \geq 0) /expansion</td>
<td>(gap_{i,t} &gt; 0) /expansion/associated with banking crisis</td>
<td>(gap_{i,t} &gt; 0) /expansion/preceded by credit boom</td>
</tr>
<tr>
<td>(\hat{\gamma}_0)</td>
<td>0.016</td>
<td>0.026</td>
</tr>
<tr>
<td>(0.017)</td>
<td>((0.03))</td>
<td>((0.71))</td>
</tr>
<tr>
<td>(\hat{\gamma}_{1}^{bus}) for gap &lt; 0</td>
<td>-</td>
<td>-0.005</td>
</tr>
<tr>
<td>(0.026)</td>
<td>((0.009))</td>
<td>((0.99))</td>
</tr>
<tr>
<td>and in recession</td>
<td>-</td>
<td>-0.004</td>
</tr>
<tr>
<td>(0.19)</td>
<td>((0.019))</td>
<td>((0.82))</td>
</tr>
<tr>
<td>(\hat{\gamma}_{1}^{fin})</td>
<td>-</td>
<td>-0.016</td>
</tr>
<tr>
<td>(0.024)</td>
<td>((0.004))</td>
<td>((0.55))</td>
</tr>
<tr>
<td>(\sum_{j=1}^{k} \hat{\delta})</td>
<td>0.006</td>
<td>0.024</td>
</tr>
<tr>
<td>(0.83)</td>
<td>(0.56)</td>
<td>(0.83)</td>
</tr>
</tbody>
</table>

Note: *, **, *** denote significance at 10, 5, and 1% levels respectively. Standard errors are in brackets, p-values in italics. The hypothesis that the sum of coefficients is equal to zero is tested by means of an F-test. The total number of observations is 1379, with \(T=92\), max \(T=93\), min \(T=86\) and \(N=15\). \(\hat{\gamma}\) depicts the slope estimate when the output gap is negative, the business cycle dummy equals 0 and the financial stress dummy equals 1.

Table 3.5 Estimation results - recessions in association with financial stress (CCEMG)

**Recoveries** Table 3.6 displays the results in case the interaction variables include the association of financial distortions with business cycle recoveries. In this case, \(\tau_{1} = (\tau_{1}^{sign}, \tau_{1}^{bus}, \tau_{1}^{fin})\) where \(\tau_{1}^{fin} = 1\) during periods of financial stress associated with recoveries. \(\hat{\gamma}\) in table 3.6 refers to the estimates for \(\hat{\gamma}_{1} = (0, 1, 1)\), such that we can analyze the effect on the Phillips curve slope during recovery episodes with restrictive credit conditions. As columns 2 to 4 indicate, the Phillips curve slope is not significantly different for recoveries that are associated with a banking crisis, when they are preceded by a credit boom or when they are characterized as creditless.

These results therefore suggest that one cannot expect an effect of the credit cycle over and above the impact on output, even for an underemployed economy. The Phillips curve slope estimate is also not expected to be altered during periods of restrictive credit conditions, either during recessions or during recoveries.
\[
\pi_{it} = \bar{\alpha}_i + \gamma \text{gap}_{it} + \sum_{j=1}^{k} \delta_j \Delta \text{gap}_{t-j} + \lambda E_i[\pi_{i,t+4}] + \sum_{j=1}^{l} \rho_j \pi_{i,t-j} + \sum_{j=0}^{\delta} \varphi_j \text{ext}_{t-j} + \varepsilon_{i,t}
\]

**Interaction model**

<table>
<thead>
<tr>
<th></th>
<th>(1)</th>
<th>(2)</th>
<th>(3)</th>
<th>(4)</th>
</tr>
</thead>
<tbody>
<tr>
<td>(D_{gap,1}^{1a}) if: (\text{gap}_{it} \geq 0) / expansion</td>
<td>/expansion/associated with banking crisis</td>
<td>/expansion/preceded by credit boom</td>
<td>/expansion/creditless recovery</td>
<td></td>
</tr>
<tr>
<td>(\hat{\gamma}_0)</td>
<td>0.016 (0.017) 0.334</td>
<td>-0.003 (0.010) 0.799</td>
<td>0.027* (0.015) 0.079</td>
<td>-0.000 (0.014) 0.993</td>
</tr>
<tr>
<td>(\hat{\gamma}<em>1) for (\text{gap}</em>{it} &lt; 0)</td>
<td>-0.022 (0.037) 0.515</td>
<td>0.021 (0.020) 0.363</td>
<td>0.002 (0.024) 0.946</td>
<td></td>
</tr>
<tr>
<td>(\gamma_{bus}^{1})</td>
<td>-0.011 (0.016) 0.513</td>
<td>-0.013 (0.011) 0.251</td>
<td>-0.019 (0.015) 0.214</td>
<td></td>
</tr>
<tr>
<td>(\gamma_{fin}^{1})</td>
<td>0.036 (0.032) 0.261</td>
<td>0.006 (0.006) 0.302</td>
<td>0.020 (0.013) 0.199</td>
<td></td>
</tr>
<tr>
<td>(\sum_{j=1}^{2} \hat{\delta})</td>
<td>0.006 (0.017) 0.334</td>
<td>0.025 (0.032) 0.242</td>
<td>0.010 (0.062) 0.627</td>
<td>0.032 (0.123) 0.123</td>
</tr>
</tbody>
</table>

Note: *, **, *** denote significance at 10, 5, and 1% levels respectively. Standard errors are in brackets, p-values in italics. The hypothesis that the sum of coefficients is equal to zero is tested by means of an F-test. The total number of observations is 1379, with \(T=92\), max \(T=93\), min \(T=86\) and \(N=15\). \(a\) \(\hat{\gamma}\) depicts the slope estimate when the output gap is negative, the business cycle dummy equals 1 and the financial stress dummy equals 1.

**Table 3.6** Estimation results - recoveries in association with financial stress (CCEMG)

\[
\pi_{it} = \bar{\alpha}_i + \gamma \text{gap}_{it} + \sum_{j=1}^{k} \delta_j \Delta \text{gap}_{t-j} + \lambda E_i[\pi_{i,t+4}] + \sum_{j=1}^{l} \rho_j \pi_{i,t-j} + \sum_{j=0}^{\delta} \varphi_j \text{ext}_{t-j} + \varepsilon_{i,t}
\]

**Interaction model**

<table>
<thead>
<tr>
<th></th>
<th>(1)</th>
<th>(2)</th>
<th>(3)</th>
<th>(4)</th>
</tr>
</thead>
<tbody>
<tr>
<td>(D_{gap,1}^{1a}) if: (\text{gap}_{it} \geq 0)</td>
<td>(D_{gap,1}^{1a}) and (D_{ext,1}^{1a}) if: (\text{gap}<em>{it} \geq 0) and (\Delta \text{gap}</em>{it} \geq 0)</td>
<td>(\text{gap}_{it} \geq 0) and (</td>
<td>\text{gap}_{it}</td>
<td>&gt; 1.5) and (\Delta \text{gap}_{it} \geq 0)</td>
</tr>
<tr>
<td>(\hat{\gamma}_0)</td>
<td>0.016 (0.017) 0.334</td>
<td>0.028 (0.010) 0.136</td>
<td>0.013 (0.018) 0.494</td>
<td>0.011 (0.017) 0.513</td>
</tr>
<tr>
<td>(\sum_{j=1}^{2} \hat{\delta}_0)</td>
<td>0.006 (0.017) 0.334</td>
<td>-0.027 (0.493) 0.483</td>
<td>0.001 (0.947) 0.964</td>
<td>-0.002 (0.921) 0.921</td>
</tr>
<tr>
<td>(\sum_{j=1}^{2} \hat{\delta})</td>
<td>- (0.017) 0.334</td>
<td>-0.025 (0.614) 0.864</td>
<td>0.018 (0.864) 0.864</td>
<td>0.100 (0.366) 0.366</td>
</tr>
</tbody>
</table>

Note: *, **, *** denote significance at 10, 5, and 1% levels respectively. Standard errors are in brackets, p-values in italics. The hypothesis that the sum of coefficients is equal to zero is tested by means of an F-test. The total number of observations is 1379, with \(T=92\), max \(T=93\), min \(T=86\) and \(N=15\).

**Table 3.7** Estimation results - speed limit effect (CCEMG)
Speed limit effect

Finally, we examine the results on the existence of a speed limit effect during periods of spare capacity. We therefore consider $\Delta gap$ interactions next to the interaction term concerning the sign of the level of the output gap variable ($\gamma_1 = \gamma_1^{\text{sign}}$). The first column in table 3.7 shows the results when the $\Delta gap$ terms are interacted with a binary dummy that equals 1 when the output gap is closing for a given level of capacity utilization (i.e., $\Delta gap_{i,t} \geq 0$). Columns 2 and 3 next display the results when the speed limit dummy is 1 only for periods of respectively large gaps and persistently large gaps. Only in the last column, the speed limit effects appear to be considerably different in case the output gap closes ($\sum_{j=1}^{2} \delta_{j} = 0.10$) relative to the other periods ($\sum_{j=1}^{2} \delta_{j} = -0.00$). The interactions are however insignificant, as is the case for the other columns. These findings suggest that speed limit effects do not arise when the aggregate economy revives, even in case of a period of persistent economic slack.

5 Robustness checks

As a first test of the robustness of the results, we examine the consequences of controlling for a time-varying trend in the level of inflation over time for the tests of asymmetries and nonlinearities depending on the sign, level and persistence of the output gap and the estimates of the speed limit effect. Table 3.8 displays the coefficient estimates when inflation and the next-year inflation expectations enter the regressions in deviation from the long-term inflation expectations. The results show that the coefficients for this inflation gap specification are highly similar to the ones in table 3.3 although the HNKPC slope is not anymore significant when there is no control for the asymmetry with respect to the sign of the output gap.

Table 3.9 depicts the results of the analysis of the effect of credit evolutions on the Phillips curve slope coefficient for the inflation gap specification and an alternative business cycle definition. We employ the classical methodology used in Claessens et al (2012) on the log of real GDP to obtain an alternative dating of the business cycle phases. The first 2 columns confirm the documented insignificance of the credit cycle on the slope coefficient in table 3.4. This also holds when the alternative business cycle dating is employed (columns 3 and 4). The credit cycle interaction is however significant at the 10
Inflation gap specification

<table>
<thead>
<tr>
<th>Linear model</th>
<th>Interaction model</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>(1)</td>
</tr>
<tr>
<td>( D_{t}^{\text{gap}} = 1 )</td>
<td>( \gamma_{0} = 0.009 )</td>
</tr>
<tr>
<td>\text{if:} ( \gamma_{t} \geq 0 )</td>
<td>( -0.039 )</td>
</tr>
<tr>
<td>( \gamma_{t} &lt; -1.5 )</td>
<td>( -0.066 )</td>
</tr>
<tr>
<td>( \gamma_{t} &gt; 1.5 )</td>
<td>( )</td>
</tr>
<tr>
<td>( \gamma_{t} )</td>
<td>( 0.061^{*} )</td>
</tr>
<tr>
<td>( \gamma_{1} = 0.053^{*} )</td>
<td>( 0.083 )</td>
</tr>
<tr>
<td>( \gamma_{1}^{\text{sign}} )</td>
<td>( 0.052 )</td>
</tr>
<tr>
<td>( \gamma_{1}^{\text{LN}} / \gamma_{1}^{\text{LP}} )</td>
<td>( 0.024 )</td>
</tr>
<tr>
<td>( \gamma_{1}^{\text{slack}} / \gamma_{1}^{\text{boom}} )</td>
<td>( 0.103^{**} )</td>
</tr>
<tr>
<td>( \sum_{j=1}^{2} \delta_{j} )</td>
<td>( 0.006 )</td>
</tr>
<tr>
<td>( \sum_{j=1}^{2} \tilde{\delta}_{j} )</td>
<td>( 0.005 )</td>
</tr>
<tr>
<td>( \sum_{j=1}^{2} \tilde{\delta}_{j} )</td>
<td>( 0.006 )</td>
</tr>
</tbody>
</table>

Note: *, **, *** denote significance at 10, 5, and 1% levels respectively. Standard errors are in brackets, p-values in italics. The hypothesis that the sum of coefficients is equal to zero is tested by means of an F-test.

The total number of observations is 1379, with \( T=92 \), max \( T=93 \), min \( T=86 \) and \( N=15 \).

Table 3.8 Robustness check - level and persistence of output gap and speed limit effect (CCEMG)
percent level and with the a-priori expected sign in case the inflation gap specification is run with the alternative business cycle dummy. The effect is however relatively small and does not lead to a HNKPC slope that is significantly different from zero. From this, we can infer that the results in table 3.4 are robust to the inflation gap specification and the alternative business cycle definition.

In table 3.10, we inspect the robustness of the estimates when recessions are analyzed in relation with the indicators of financial stress episodes. When the alternative business cycle is employed, we confine the attention to the credit cycle since the amount of economic contractions associated with credit booms is too limited according to this criterion. Overall, the message remains the same than in table 3.5, the credit cycle does not seem to affect the Phillips curve.

Finally, it is clear from table 3.11 that also the results for financial stress periods associated with recoveries are fairly similar to table 3.6. Creditless recoveries now however exert a significant negative effect on the HNKPC slope when the inflation gap specification is employed (column 3). The eventual $\gamma$ estimate however remains insignificant from zero.

In conclusion, the previous findings are robust to the use of the inflation gap specification and the use of an alternative business cycle dating method.
Inflation gap specification & alternative business cycle definition

<table>
<thead>
<tr>
<th>Specification</th>
<th>(1)</th>
<th>(2)</th>
<th>(3)</th>
<th>(4)</th>
<th>(5)</th>
<th>(6)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Upturns</td>
<td>0.009</td>
<td>-0.005</td>
<td>-0.006</td>
<td>-0.01</td>
<td>-0.007</td>
<td>-0.003</td>
</tr>
<tr>
<td>Credit booms</td>
<td>(0.002)</td>
<td>(0.002)</td>
<td>(0.002)</td>
<td>(0.002)</td>
<td>(0.002)</td>
<td>(0.002)</td>
</tr>
</tbody>
</table>

Note: *, **, *** denote significance at 10, 5, and 1% levels respectively. Standard errors are in brackets, p-values in italics. The hypothesis that the sum of coefficients is equal to zero is tested by means of an F-test. The total number of observations is 1379, with $T=92$, $\min T=86$ and $N=15$.

\( \gamma \) depicts the slope estimate when the output gap is negative.
Table 3.10 Robustness check - recessions in association with financial stress (CCEMG)

<table>
<thead>
<tr>
<th></th>
<th>Inflation gap specification</th>
<th>Alternative business cycle definition</th>
<th>Inflation gap specification &amp; alternative business cycle definition</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>(1)</td>
<td>(2)</td>
<td>(3)</td>
</tr>
<tr>
<td></td>
<td>associated with banking crisis</td>
<td>preceded by credit boom</td>
<td>associated with banking crisis</td>
</tr>
<tr>
<td>$\hat{\gamma}_0$</td>
<td>0.006 (0.032) 0.840</td>
<td>0.016 (0.022) 0.474</td>
<td>-0.023 (0.028) 0.410</td>
</tr>
<tr>
<td>$\hat{\gamma}$ for gap &lt; 0 and in recessions</td>
<td>-0.016 (0.041) 0.691</td>
<td>0.011 (0.029) 0.700</td>
<td>-0.030 (0.034) 0.372</td>
</tr>
<tr>
<td>$\hat{\gamma}_{bus}$</td>
<td>0.003 (0.019) 0.893</td>
<td>-0.009 (0.016) 0.571</td>
<td>0.022 (0.020) 0.280</td>
</tr>
<tr>
<td>$\hat{\gamma}_{fin}$</td>
<td>-0.023 (0.025) 0.368</td>
<td>-0.004 (0.019) 0.822</td>
<td>-0.007 (0.019) 0.702</td>
</tr>
<tr>
<td>$\sum_{j=1}^{2} \tilde{\delta}$</td>
<td>0.008 (0.765) 0.873</td>
<td>0.004 (0.873)</td>
<td>-0.027 (0.596)</td>
</tr>
</tbody>
</table>

Note: **, ***, *** denote significance at 10, 5, and 1% levels respectively. Standard errors are in brackets, p-values in italics. The hypothesis that the sum of $\delta$ coefficients is equal to zero is tested by means of an F-test. The total number of observations is 1379, with $T=92$, max $T=93$, min $T=86$ and $N=15$. $\hat{\gamma}$ depicts the slope estimate when the output gap is negative, the business cycle dummy equals 0 and the financial stress dummy equals 1.
### Table 3.11 Robustness check - recoveries in association with financial stress (CCEMG)

<table>
<thead>
<tr>
<th></th>
<th>Inflation gap specification</th>
<th>Alternative business cycle definition</th>
<th>Inflation gap specification &amp; alternative business cycle definition</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>(1)</td>
<td>(2)</td>
<td>(3)</td>
</tr>
<tr>
<td>$\hat{\gamma}_0$</td>
<td>$-0.007$ (0.015) 0.645</td>
<td>$0.014$ (0.014) 0.329</td>
<td>$-0.004$ (0.020) 0.838</td>
</tr>
<tr>
<td>$\hat{\gamma}_{gap&lt;0}$</td>
<td>$0.029$ (0.051) 0.563</td>
<td>$-0.002$ (0.021) 0.914</td>
<td>$0.003$ (0.030) 0.931</td>
</tr>
<tr>
<td>$\hat{\gamma}_{bus}$</td>
<td>$-0.024$ (0.020) 0.221</td>
<td>$-0.021$ (0.018) 0.165</td>
<td>$-0.029^*$ (0.015) 0.050</td>
</tr>
<tr>
<td>$\hat{\gamma}_{fin}$</td>
<td>$0.060$ (0.044) 0.170</td>
<td>$0.005$ (0.004) 0.259</td>
<td>$-0.036^{**}$ (0.016) 0.923</td>
</tr>
<tr>
<td>$\sum_{j=1}^{2} \hat{\delta}$</td>
<td>$0.023$ (0.328) 0.318</td>
<td>$0.013$ (0.334) 0.356</td>
<td>$0.041$ (0.017) 0.017</td>
</tr>
</tbody>
</table>

Note: **, *** denote significance at 10, 5, and 1% levels respectively. Standard errors are in brackets, p-values in italics. The hypothesis that the sum of $\delta$ coefficients is equal to zero is tested by means of an F-test. The total number of observations is 1379, with $T=92$, max $T=93$, min $T=86$ and $N=15$.

$\hat{\gamma}$ depicts the slope estimate when the output gap is negative and when both the business cycle dummy and the financial stress dummy equal 1.
6 Conclusions

In light of the global financial crisis of 2008-2009 and the hypothesized missing disinflation puzzle, we estimate a hybrid New Keynesian Phillips curve for a panel of 15 advanced OECD economies while considering the evolution of bank credit to the private sector and the occurrence of financial stress periods next to possible nonlinearities and asymmetries depending on the level of economic activity. The goal is to examine why the reaction of inflation has been relatively mild in the immediate aftermath of the global crisis and whether restrictive credit conditions have lowered this reaction. Overall, we find that the Phillips curve is "alive". The output gap matters, at least when the economy is booming. Our findings suggest that the Phillips curve is significantly steeper for positive output gaps relative to negative output gaps and that the slope is insignificant for the latter. The magnitude of deviations of GDP from potential does not have an additional influence on the reaction of inflation once one controls for the sign of the deviation. Also persistently large output gaps do not exert an influence on the Phillips curve slope. Negative output gaps were omnipresent since 2009 and the fairly stable inflation rates and the non-realized deflation fears subsequent to the financial crisis are thus consistent with the established flat Phillips curve during these episodes.

Given the significant asymmetry in the Phillips curve with respect to the sign of the output gap, we accordingly focus on episodes of spare production capacity to examine the Phillips curve relationship in relation with bank credit evolutions. We find no evidence for different inflation dynamics driven by the credit cycle. The reaction of inflation to the output gap is not attenuated during credit cycle downturns relative to credit booms and more general credit upturns. We also do not find evidence of an effect on the slope of the Phillips curve during recoveries and recessions associated with periods of financial stress. We therefore conclude that there is no firm evidence in our sample that bank credit evolutions and financial distress have significantly altered the Phillips curve relationship following the global financial crisis. The mild inflation reaction subsequent to the financial crisis can thus be linked to an asymmetric reaction towards the extent of spare production capacity whereas the occurrence of a credit downturn for 13 of the 15 countries occurring in association with the global financial crisis does not seem to have counteracted the expectations of disinflation.

We further have no evidence that underpins the existence of a speed limit effect on
inflation when the extent of spare capacity shrinks. The speed limit effect is found to be
insignificant, even in presence of persistent and severe economic slack periods. The fact
that a long-lasting and pronounced contraction is more likely to alter production resources
obsolete and inadequate does not seem to generate additional inflationary pressures.
Appendices

A. Data sources

Core inflation: consumer price index, all items non-food, non-energy, growth rates with respect to previous quarter, seasonally adjusted [OECD, Main Economic Indicators].

Nominal/Real GDP: gross domestic product in billions of national currency, value/volume, market prices [OECD, Economic Outlook No 93].

Output gap: output gap of the total economy [OECD, Economic Outlook No 90 (for historical episodes) and 93 (for post-2010 episodes)].

One-year-ahead inflation expectations: expectations of headline inflation for the next year, annualized growth with respect to the previous year, transformed to quarterly growth rates [Consensus Economics].

Long-term inflation expectations: expectations of headline inflation for the next 6 to 10 years, annualized growth with respect to the previous year, transformed to quarterly growth rates [Consensus Economics]. Observations on the 6-to-10-years inflation expectations are however more limited in time and over countries. Missing observations from 1990Q1 onwards are replaced by a HP-filtered trend on headline CPI inflation since 1960Q1 of the first available quarter to obtain balanced series over the entire sample period.\(^{22}\)

Commodity imports price inflation: price of commodity imports, growth rates with respect to previous quarter, seasonally adjusted [OECD, Economic Outlook No 93].

Non-commodity imports of goods and services price inflation: price of non-commodity imports of goods and services, growth rates with respect to previous quarter, seasonally adjusted [OECD, Economic Outlook No 93].

\(^{22}\)There are no data on these long-term inflation expectations for Denmark, Finland and Ireland. For Australia, South Korea, the Netherlands, Norway and Sweden, the survey data series start in, respectively, 1991Q2, 1995Q2, 1995Q2, 1998Q4 and 1995Q2. The data are collected from 1990Q2 onwards and the 1990Q1 observations are thus based on the HP trend for all countries.
<table>
<thead>
<tr>
<th>COUNTRY</th>
<th>TIME SPAN</th>
<th>NUMBER</th>
<th>COUNTRY</th>
<th>TIME SPAN</th>
<th>NUMBER</th>
</tr>
</thead>
<tbody>
<tr>
<td>Australia</td>
<td>1990Q1</td>
<td>2013q1</td>
<td>93</td>
<td>Japan</td>
<td>1990Q1</td>
</tr>
<tr>
<td>Canada</td>
<td>1990Q1</td>
<td>2013q1</td>
<td>93</td>
<td>Korea, South</td>
<td>1991Q2</td>
</tr>
<tr>
<td>Denmark</td>
<td>1990Q1</td>
<td>2013q1</td>
<td>93</td>
<td>Netherlands</td>
<td>1990Q1</td>
</tr>
<tr>
<td>Finland</td>
<td>1990Q1</td>
<td>2013q1</td>
<td>93</td>
<td>Norway</td>
<td>1990Q1</td>
</tr>
<tr>
<td>France</td>
<td>1990Q1</td>
<td>2013q1</td>
<td>93</td>
<td>Sweden</td>
<td>1990Q1</td>
</tr>
<tr>
<td>Germany</td>
<td>1991Q4</td>
<td>2013q1</td>
<td>86</td>
<td>UK</td>
<td>1990Q1</td>
</tr>
<tr>
<td>Ireland</td>
<td>1991Q1</td>
<td>2013q1</td>
<td>89</td>
<td>US</td>
<td>1990Q1</td>
</tr>
<tr>
<td>Italy</td>
<td>1990Q1</td>
<td>2013q1</td>
<td>93</td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

Table 3.12 Overview number of observations per country (for baseline model in (1))

**Total bank credit to the private non-financial sector:** long series on total credit and domestic bank credit to the private nonfinancial sector in billions of national currency, adjusted for breaks, annualized [BIS].

**Headline CPI:** consumer price index, all items, index 2010=100 [OECD, Main Economic Indicators].

**Population:** total population, midyear estimates [World bank, World Development Indicators database].

**B. Sample coverage**

**C. Cycle definitions**

**Persistent economic boom or slack:** periods with output gap >1.5 or <1.5 for at least 8 consecutive quarters. This definition is based on the employed threshold value in Meier (2010) for economic slack periods and the estimated threshold values in Barnes and Olivei (2003) for the unemployment gap of the US.\(^23\)

**Business cycle:** recessions and expansions in economic activity are defined based on the output gap measure. Expansions run from trough to peak (excluding the trough)

\(^23\)Note that the cycle datings are based on data since 1960Q1 or the first available quarter, particular episodes starting before 1990Q1 but continuing through the sample under consideration can be accordingly defined such that the number of observations is maximized.
and recessions run from peak to trough (excluding the peak). A trough is identified as a period in which the output gap is below minus its standard deviation. To avoid double dip episodes, it is assumed that there are at least 8 quarters between two troughs. If several troughs occur within 8 quarters, the quarter with the lowest value of the output gap is considered to be the trough period. A peak is subsequently defined as the quarter with the largest output gap value between 2 consecutive troughs. It is further imposed that the recessions and expansions last for at least 2 quarters. If a business cycle phase is defined to last for only 1 quarter, this period is considered to belong to the other phase. This methodology is based on the work of Sugawara and Zalduendo (2013) although they employ Hodrick-Prescott (HP) filtering on real GDP to infer the cyclical variation in output whereas we make use of OECD output gap series where potential output is estimated based on a multivariate production function methodology. The first 8 quarters of the expansions are defined as recovery periods.

**Banking crises:** we use the (quarterly) dating of Drehmann, Borio and Tsatsaronis (2011) and supplement it with the (annual) dating of Laeven and Valencia (2013).

**Credit cycle:** we employ the classical methodology used in Claessens et al (2012) on the real credit data to define upturns and downturns in bank credit per capita. As in Claessens et al (2012), the algorithm of Harding and Pagan (2002) is applied on the log level of real credit to determine the cycle’s turning points. The value $y_t$ is considered as a peak at time $t$, if $(y_t - y_{t-2}) > 0$, $(y_t - y_{t-1}) > 0$, $(y_{t+2} - y_t) < 0$ and $(y_{t+1} - y_t) < 0$, while a trough occurs at time $t$ if $(y_t - y_{t-2}) < 0$, $(y_t - y_{t-1}) < 0$, $(y_{t+2} - y_t) > 0$ and $(y_{t+1} - y_t) > 0$. The downturn then runs from peak to trough and the upturn from trough to the next peak. The restriction is thereby imposed that the downturn lasts at least 2 quarters and the upturn at least 5 quarters.

**Credit booms:** similar to Mendoza and Terrones (2008), a HP filter on the log of real credit per capita from 1960Q1 (or from the first available observation) onwards is applied to obtain deviations of credit from its country-specific long-run trend. Boom periods are defined as consecutive quarters for which the cyclical component is equal or larger than 1.5 times its standard deviation. Credit data are deflated by headline CPI. The peak of a credit boom is subsequently defined as the quarter from the set of consecutive quarters that satisfy the credit boom condition with the largest
difference between the cyclical component and its standard deviation. Given the peak, the start date of the boom is the quarter before the peak with the minimum difference between the cyclical component and its standard deviation. The end date of the boom is similarly determined as the quarter after the peak that satisfies this latter condition.

**Creditless recoveries:** a creditless recovery is defined as a recovery without a pick up in real credit per capita. Based on the definition in the work of Sugawara and Zalduendo (2008), the recovery is considered to be creditless when the average growth rate of real credit per capita (deflated by headline CPI) during the recovery is less then or equal to zero.

**Alternative business cycle definition:** as in Claessens et al (2012), the algorithm of Harding and Pagan (2002) is applied on the log level of real GDP to define recessions and expansions in the business cycle. A recession runs from peak to trough and the upturn from trough to the next peak. Again, the restriction is imposed that the recession lasts at least 2 quarters and the upturn at least 5 quarters.
Bibliography


CHAPTER 4

Wage Indexation and the Monetary Policy Regime

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Ghent University  Ghent University  Ghent University

Abstract

We estimate a New Keynesian wage Phillips curve for a panel of 24 OECD countries, and allow the degree of wage indexation to past inflation to vary according to the monetary policy regime. We find that the extent of wage indexation is significantly lower in an inflation targeting regime, in contrast to monetary targeting, exchange rate targeting and policy regimes without an explicit quantitative anchor. The results put into question whether embedding a constant degree of wage indexation in standard DSGE models is truly structural.

*JEL classification:* C23, E42, J30

*Keywords:* wage indexation, regimes, cross-country panel, Phillips curve

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1 Introduction

New Keynesian dynamic stochastic general equilibrium (DSGE) models typically assume sticky wages and partial wage indexation to past inflation. Notably, the degree to which wages are indexed to past inflation is hard-wired as a fixed and policy invariant parameter (e.g. Christiano, Eichenbaum and Evans, 2005; Smets and Wouters, 2007). The assumption of a constant degree of wage indexation, however, has been rejected by institutional and empirical evidence for the United States (US). In particular, Holland (1986) documents a substantial rise in the proportion of wage contracts with indexation clauses to price level changes in the US between the late 1960s and mid 1980s, after which there was again a decline. Hofmann, Peersman and Straub (2012) estimate the extent of wage indexation in the US over time, and find a considerably higher degree of indexation during the "Great Inflation" of the 1970s compared to the earlier and later periods. Holland (1986) attributes the rise of indexation practices in the 1970s to much higher inflation uncertainty, whereas Hofmann et al. (2012) explain the rise and fall of wage indexation by a weaker reaction of the Federal Reserve to inflation during the "Great Inflation", and more aggressive inflation stabilization before and after this period. Specifically, a weakly inflation stabilizing monetary policy regime is conducive to high and volatile inflation, which fosters the use of wage indexation clauses as protection against inflation uncertainty.

To the extent that inflation uncertainty is determined by the monetary policy regime, a possible link between the degree of wage indexation and monetary policy is supported by economic theory. Gray (1978), for example, shows in a neoclassical model with wage rigidities that the optimal proportion of wage contracts indexed to inflation increases with the variance of monetary disturbances. Ehrenberg, Danziger and San (1983) further show in an efficient contract model that the gain of indexation for risk averse workers, and hence the likelihood of indexation, rises when inflation uncertainty is higher. On the other hand, Carrillo, Peersman and Wauters (2014) demonstrate that utility maximizing workers only want to index wages to past inflation when permanent shocks to the inflation target (and technology) dominate output fluctuations, but not when temporal inflation target (and aggregate demand) shocks dominate.

In this paper, we formally examine whether wage indexation varies across monetary policy regimes.\footnote{Although empirical work has found that higher inflation uncertainty raises the prevalence of cost-of-} More precisely, we estimate the reduced-form empirical New Keynesian
wage Phillips curve of Galí (2011) on a panel dataset covering 24 OECD countries between 1960Q1 and 2011Q4, and allow the degree of wage indexation to vary according to the monetary policy regime. Since the monetary policy regime of an individual country is in general quite stable over time, a panel dataset approach increases the number of observations significantly, which allows us to formally estimate the role of the policy regime. We control for labor market institutions and include an estimate of the time-varying inflation target of the central bank to examine the importance of wage indexation once variation in trend inflation is taken into account.

We identify the monetary policy regime of a country in a specific period based on the presence of an explicit quantitative monetary target, which takes three forms: inflation, money growth and exchange rates targets. Quantitative targets are transparent policy indicators and can easily be measured.\(^2\) A formal commitment to a quantitative target is expected to influence (improve) the formation of inflation expectations and (reduce) inflation uncertainty of workers (Mishkin, 2007). We distinguish between the presence of an inflation, money growth and exchange rate target, because the underlying dynamics of these strategies and formation of inflation expectations are inherently different. For example, inflation targeting central banks typically try to stabilize inflation in the short to medium term, whereas money growth targeting is more a commitment to low inflation in the long run.\(^3\)

The estimation results provide a number of important considerations for macroeconomic analysis and policymakers. First, we find that wage indexation is significant and economically relevant for the sample under analysis. Second, the results confirm that wage indexation varies across monetary policy regimes. The degree of wage indexation to past inflation turns out to be significantly lower in a regime which has a quantitative inflation target but this condition does not hold for money growth or exchange rate targets. Since

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\(^2\)There is a large literature that has focused on domestic institutions to represent the monetary policy regime, such as the degree of central bank independence and indicators of transparency (e.g. Alesina and Summers, 1993; Eijffinger and Geraats, 2006; Dincer and Eichengreen, 2014). The use of such indicators, however, would reduce our panel dataset considerably. Moreover, the construction of these indicators involves debatable personal judgment.

\(^3\)The inflationary outcomes of the three different types of nominal anchors also turn out to be different (Fatás, Mihov and Rose, 2007).

---
the extent of wage indexation is different across monetary policy regimes, the constant
indexation assumption embedded in standard DSGE models is susceptible to the Lucas
(1976) critique, i.e. it is not intrinsic to the deep structure of the economy and not a policy
invariant parameter. The analysis of alterations to the policy regime and counterfactual
policy simulations in these models are thus potentially misleading. Similarly, the compu-
tation of optimal monetary policies that are conditional on the estimated parameters of
these models may produce unreliable results if the optimal policy strategy implies changes
to the degree of wage indexation and thus the structure of the economy.

Third, to the extent that having a quantitative inflation target itself is the key mech-
anism that lowers wage indexation to past inflation, the results suggest that the adoption
of an explicit inflation target could reduce the inflationary consequences of shocks hitting
the economy, the costs of disinflation, and the volatility of output and prices. Specifically,
given that inflation is driven by real marginal costs, which are directly linked to
wages, a significant reduction in the degree of wage indexation implies less amplification
of the inflationary consequences of shocks hitting the economy due to mutually reinforc-
ing feedback effects between wages and prices, i.e. less so-called second-round effects of
inflationary shocks. A reduction in the inflationary effects of shocks hitting the economy
requires in turn less aggressive monetary policy responses to stabilize inflation, lowering
also output volatility (Hofmann et al., 2012).

Our work is related to several strands in the literature. Messina and Sanz-de-Galdeano
(2014) use micro level data to document how Brazil’s and Uruguay’s disinflation policies
changed the nature of wage rigidities. For Brazil in particular, they document how the
introduction of inflation targeting affected wage indexation. Alogoskoufis and Smith (1991)
study wage and price inflation series from 1892 to 1987 for the US and the UK; they report
coinciding shifts in the wage Phillips curve and price inflation persistence, which they
link to departures from international fixed exchange rate regimes. Muto and Shintani
(2014) perform an empirical evaluation of the New Keynesian Wage Phillips Curve for
Japan and the US. They show with rolling window regressions that the importance of

\footnote{Based on the estimations, we can only conclude that the degree of wage indexation is different across
monetary policy regimes, and significantly lower in regimes that have a quantitative inflation target. Whether
the inflation target itself is the mechanism that reduces indexation practices is a question out of
the scope of this paper. Specifically, other features of a monetary policy regime that are typically asso-
ciated with an explicit inflation target could reduce the degree of wage indexation, for instance enhanced
transparency, more independent central banks or an inflation averse society.}
wage indexation has declined over time for both countries, which they link to lower and more stable inflation. Benati (2008) questions whether the intrinsic inflation persistence found in post-WWII US data is truly structural. He estimates the price Phillips curve on historical data for a set of countries and finds the price indexation parameter to be very low or zero under stable monetary policy regimes with clearly defined nominal anchors. Levin, Natalucci and Piger (2004) find that inflation expectations appear to be more forward looking, and inflation less persistent, in inflation targeting countries. A related study is also Fatás et al. (2007), who find that having an explicit quantitative target for monetary policy, in particular an inflation target, is systematically related to a lower average level of inflation. Finally, our study is related to the literature that analyzes the role of monetary policy institutions for inflation outcomes and economic growth, such as central bank independence (Alesina and Summers, 1993) and transparency (Sterne, Stasavage and Chortareas, 2002; Eijffinger and Geraats, 2006; Dincer and Eichengreen, 2014).

The remainder of the paper is organized as follows: In the next section, we present the estimation results for a benchmark wage Phillips curve model with a constant degree of wage indexation. In section 3, we extend the benchmark model to analyze the influence of the monetary policy regime on the extent of backward-looking wage indexation, controlling for a set of labor market characteristics. Finally, section 4 concludes.

2 Wage Phillips curve with constant indexation

In this section, we present the estimation results of a wage Phillips curve, assuming a constant and policy invariant degree of wage indexation. We first derive a benchmark empirical New Keynesian wage Phillips curve in section 2.1. Section 2.2 presents the data and discusses some econometric issues, whereas the estimation results are shown in section 2.3.

2.1 Model specification

Theoretical framework Our theoretical framework is drawn from Galí (2011), who derives the empirical wage Phillips curve from a New Keynesian model that includes the unemployment rate. He thereby provides both a theoretical foundation for the empirical relation and a structural interpretation of the reduced form coefficients. The model
assumes staggered wage setting as in Erceg, Henderson and Levin (2000), which means that a worker’s wage cannot be re-optimized in every period. When the wage cannot be reset, it is assumed to index to a weighted average of past price inflation ($\pi_{t-1}^p$) and the central bank’s inflation target ($\pi^*$), and additionally trend productivity growth ($g$):

$$\gamma\pi_{t-1}^p + (1-\gamma)\pi^* + g,$$

with $\gamma \in (0, 1)$ as the weight to past inflation. We denote wage inflation by $\pi_t^w$ and the difference between unemployment and the natural rate by $\hat{u}_t \equiv u_t - u^n$.

The model’s solution is given by

$$\pi_t^w = \alpha + \gamma\pi_{t-1}^p + \psi_0\hat{u}_t + \psi_1\hat{u}_{t-1},$$

where $\alpha \equiv (1 - \gamma)\pi^* + g$ (Galí, 2011, equation 19).

**Econometric model** Bringing (4.1) to a panel data setting results in the econometric *benchmark wage inflation model*

$$\pi_{i,t}^w = \alpha_i + \gamma\pi_{i,t-1}^p + \psi_{0i}\hat{u}_{i,t} + \psi_{1i}\hat{u}_{i,t-1} + \eta_{i,t},$$

with subscripts $i$ and $t$ denoting the country and time period and $\alpha_i$ referring to $((1 - \gamma)\pi^* + g_i - (\psi_0 + \psi_1)u^n_i$ and other time-invariant effects. Our main interest is the degree of indexation to past price inflation ($\gamma$), which is expected to lie between 0 and 1.

As a robustness check, we extend the benchmark wage inflation model (4.2) by including a time-varying inflation target of the central bank ($\pi^*_t$), as Cogley and Sbordone (2008) argue that the price indexation parameter in a standard DSGE becomes zero once a time-varying trend inflation of the central bank is introduced. More specifically, for each country, we estimate trend inflation ($\pi^*_t$) with the AR-Trend-bound model of Chan, Koop and Potter (2013) as

$$\pi_t^p - \pi^*_t = \rho_t (\pi_{t-1}^p - \pi^*_t) + \epsilon_t,$$

where $\epsilon_t \sim N(0, 1)$ and the central bank’s trend inflation, the persistence of the inflation gap ($\rho_t$), and the log-volatility of the error term ($h_t$) all follow a random walk. The estimation details are provided in Appendix B. As trend inflation is time-varying, it no longer appears in the intercept of the benchmark wage inflation model (4.2). Accordingly, specification (4.2) is extended to the *benchmark wage gap model form*5:

$$\tilde{\pi}_{i,t}^w = \alpha_i + \gamma\tilde{\pi}_{i,t-1}^p + \psi_{0i}\hat{u}_{i,t} + \psi_{1i}\hat{u}_{i,t-1} + \eta_{i,t},$$

5Equation (4.4) is obtained under the theoretical assumption that the coefficients on past and trend inflation ($\gamma$ and $\gamma^*$) in $\pi_{i,t}^w = \alpha_i + \gamma\pi_{i,t-1}^p + \gamma^*\pi_{i,t-1}^* + \psi_{0i}\hat{u}_{i,t} + \psi_{1i}\hat{u}_{i,t-1} + \eta_{i,t}$ sum to one. A Wald test on the estimated panel coefficients did not reject this assumption (infra, p.120).
where $\pi_t^w \equiv \pi_t^w - \pi_t^*$ and $\pi_{t-1}^p \equiv \pi_{t-1}^p - \pi_t^*$.

Note that we make the implicit assumption in (4.4) that there is no immediate effect of $\pi_t^w$ on $\pi_t^*$. Given that our measure of trend inflation captures inflation expectations at an infinite horizon, this assumption is very likely to hold in practice.

### 2.2 Panel dataset and econometric considerations

We use an unbalanced panel covering quarterly data between 1960Q1 and 2011Q4 for 24 OECD economies.\(^6\) In general, an individual country’s monetary policy regime is quite stable over time and in some cases, a regime can even last for decades. We therefore use the information from a group of countries to broaden the information set and to increase the power of the test whether the degree of wage indexation depends on the type of monetary policy regime. Our wage index consists of the average hourly earnings of employees in the manufacturing sector, sourced from the OECD MEI database. The earnings data are comparable to wage rate series that proxy for the basic wages or cost-of-living allowances, but they provide a more complete measure of the overall wage income because they also include premium pay for overtime and bonuses.\(^7\) Our price measure consists of the quarterly all-items consumer price index. We construct quarter-on-quarter wage and price inflation series as 100 times the log differences of the wage and price indices. We follow Galí (2011) in taking the average of the 4-quarter lags of past inflation as a smoothed price indexation measure. The unemployment rate is expressed as a percentage of the total labor force. Appendix A provides further details on the coverage and data definitions.

There are two issues of panel estimation that we need to take into account. First, we have to verify the appropriateness of homogenous regression parameters. When the regression does not contain lagged dependent variables, as in equation (4.2), and the estimators are strictly exogenous, both homogeneous and heterogeneous estimators deliver consistent coefficient estimates (Pesaran and Smith, 1995). We nevertheless check whether

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\(^6\)The 24 countries are: Australia, Austria, Belgium, Canada, Czech Republic, Denmark, Estonia, Finland, France, Germany, Hungary, Ireland, Italy, South Korea, Netherlands, New Zealand, Norway, Poland, Portugal, Slovak Republic, Spain, Sweden, United Kingdom and the United States.

\(^7\)Compensation rates, which also include employer contributions to social security or social insurance schemes, are a widely used alternative measure. However, this series is unavailable for 10 of the 24 countries in our sample.
the assumption of homogeneous parameters affects our small sample coefficient estimates by comparing Fixed Effects (FE) and Mean Group (MG) estimation results with a Wald test.

Second, it is well known that macro panel estimates can be affected by cross-sectional dependence in the errors. Such dependence generally indicates the influence of factors that are common across countries but not explicitly modeled. To ensure consistent estimates, we apply the cross section dependence test (CD test) of Pesaran (2004) to the residuals of the equations (4.2) and (4.4). If the test indicates a significant correlation between the cross section errors, we apply Common Correlated Effects (CCE) estimators.\(^8\)

### 2.3 Results

Table 4.1 depicts the estimation results for the benchmark wage inflation model (4.2) in columns 1 and 2, while the estimates for the trend inflation adjusted wage gap model (4.4) are given in columns 3 and 4.\(^9\) A test on the homogeneity assumption of the coefficients validates the use of pooled estimators.\(^10\)

The FE estimation in column 1 indicates a substantial amount of backward wage indexation. The point estimate of \(\gamma\) is 0.81, which is within the estimated range of 0.52 and 0.83 obtained by Galí (2011) for the US. The negative contemporaneous and positive lagged effect of unemployment is in line with theoretical expectations (Galí, 2011). On impact, a 1 percentage point decline in the unemployment rate leads to a 0.37% increase in nominal wages. Heteroskedasticity and autocorrelation robust standard errors are employed to obtain a consistent inference. All parameters are statistically significant at the 1% level.

The presence of a significant cross-sectional correlation in the residuals according to the CD test (bottom of Table 4.1) results makes a case for CCE estimators. Adding the cross

\(^8\)The CCE estimators conveniently abstract from the possible influence of common factors by augmenting the observed regressors with the cross section averages (CSAs) of all variables, leaving least squares estimation still adequate. This approach yields consistent and asymptotically normal parameter estimates in a cross-country panel with a fairly large number of countries (Pesaran, 2006).

\(^9\)The hypothesis that the estimated coefficients on past and trend inflation sum to 1 is not rejected by a Wald test (value=1.04, F-statistic=0.26, p-value=0.61).

\(^10\)We performed the empirical analysis in STATA 12 with the user-written \texttt{xtcd} and \texttt{xtmg} routines of Eberhardt (2012). Wald tests on the null of coefficient homogeneity of FE versus MG generate insignificant chi-square test statistics of 3.88 and 4.39. The Wald tests on the CCE pooled and MG estimates also cannot reject the null of homogeneity (the test statistics are 2.36 and 3.07).
## Table 4.1 Results benchmark model

<table>
<thead>
<tr>
<th>Regressand:</th>
<th>wage inflation ($\pi_{i,t}^{u}$)</th>
<th>wage gap ($\pi_{i,t}^{u}$)</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>(1)</td>
<td>(2)</td>
</tr>
<tr>
<td>FE</td>
<td></td>
<td></td>
</tr>
<tr>
<td>CCEP</td>
<td></td>
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<table>
<thead>
<tr>
<th></th>
<th>FE</th>
<th>CCEP</th>
<th>FE</th>
<th>CCEP</th>
</tr>
</thead>
<tbody>
<tr>
<td>Unemployment rate ($u_t$)</td>
<td>-0.374***</td>
<td>-0.195*</td>
<td>-0.381***</td>
<td>-0.168**</td>
</tr>
<tr>
<td></td>
<td>(0.089)</td>
<td>(0.097)</td>
<td>(0.082)</td>
<td>(0.085)</td>
</tr>
<tr>
<td>Lagged unemployment ($u_{t-1}$)</td>
<td>0.269***</td>
<td>0.108</td>
<td>0.277**</td>
<td>0.125</td>
</tr>
<tr>
<td></td>
<td>(0.090)</td>
<td>(0.119)</td>
<td>(0.109)</td>
<td>(0.085)</td>
</tr>
<tr>
<td>Lagged price inflation ($\pi_{i,t}^{p}$)</td>
<td>0.806***</td>
<td>0.408***</td>
<td>-</td>
<td>-</td>
</tr>
<tr>
<td></td>
<td>(0.043)</td>
<td>(0.091)</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Lagged price gap ($\tilde{\pi}_{i,t-1}^{p}$)</td>
<td>-</td>
<td>-</td>
<td>0.285***</td>
<td>0.456***</td>
</tr>
<tr>
<td></td>
<td></td>
<td></td>
<td>(0.090)</td>
<td>(0.060)</td>
</tr>
</tbody>
</table>

| CD test (average correlation) | 0.090 | -0.035 | 0.053 | -0.042 |
| CD test (statistic and p-value) | 15.95 | -5.73  | 9.77  | -6.88  |

Note: *, **, *** denote significance at 10, 5, and 1% levels respectively.

Robust standard errors are in brackets.

Sample: $T=141$, max $T=207$, min $T=47$ and $N=24$

---

section averages (CSAs) as regressors decreases the average residual correlation although it remains significant (column 2). All coefficients decline to around half their previous size, with the indexation coefficient now attaining 0.41.

Compared to the constant trend inflation estimates in columns 1 and 2, the coefficients of the wage gap specification with time-varying trend inflation in columns 3 and 4 are broadly similar, although the FE indexation parameter shrinks somewhat. The FE and CCEP estimators furthermore result in a less pronounced decline in the extent of cross-sectional residual correlation. Interestingly, wage indexation is still statistically significant and economically important, with values of respectively 0.29 and 0.46 for the FE and CCEP estimators. This result contrasts with Cogley and Sbordone (2008), who estimate a DSGE model which incorporates drifts in trend inflation on US data and find price indexation to be essentially zero. Our approach differs from theirs in the sense that

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11By using conventional OLS variance estimates, trend inflation is treated as known and the potential generated regressor bias on the variance estimates (Pagan, 1984) is ignored. A two-step bootstrapping approach to compute the standard errors is however infeasible given the stochastic volatility-model to estimate trend inflation. We therefore rely on robust standard errors that are consistent in the presence of unknown forms of autocorrelation and heteroskedasticity.

12Cogley and Sbordone (2008) extend the Calvo pricing model in a standard DSGE model to incorporate
we assume full indexation of wages to a weighted average of past and trend inflation, we focus on the dynamics of wages instead of prices, and our empirical approach concerns a single-equation model for a panel of 24 countries.

Given that the results of both models (4.2) and (4.4) do not signal important differences concerning the significance of the lagged inflation variable, we confine the attention to the extensions of the benchmark wage inflation model in the remainder of this work.

3 Phillips curve with variable wage indexation

3.1 Wage indexation and the monetary policy regime

Motivation

A number of theoretical studies conclude that the degree of wage indexation to past inflation may depend on inflation uncertainty and the conduct of monetary policy. Specifically, Gray (1978) presents a neoclassical model with short-term wage rigidities and uncertainty, and shows that the degree of wage indexation that minimizes the deviation of output from full-information output increases with the variability of monetary disturbances. Ehrenberg et al. (1983) demonstrate in an efficient contract model that a rise in inflation uncertainty may lead to greater use of wage indexation because wage indexation helps to insulate the worker’s real wage from the effects of unanticipated inflation, whereas it reduces the impact of lower-than-anticipated inflation on the real cost of labor inputs for firms. Carrillo et al. (2014) show that utility maximizing wage setters raise the extent of wage indexation to past inflation when the variability of permanent shocks to the inflation target of the central bank increases, whereas the amount of indexation declines when there is a rise in the volatility of temporal inflation target shocks.

A possible link between wage indexation, inflation uncertainty and monetary policy is also supported by institutional and empirical evidence for the US. Holland (1986) demonstrates drifts in trend inflation and derive a reduced form New Keynesian Phillips Curve (NKPC) with time-varying coefficients. This reduced form NKPC is estimated on US data by means of a Bayesian time-varying VAR. Under the assumption that non-optimized prices are fully indexed to a mixture of current trend inflation and one-period lagged inflation, the reduced form NKPC collapses to a more traditional NKPC based on the inflation gap with constant coefficients and without additional forward-looking terms. This latter specification is similar to our gap specification.
Chapter 4

strates that the proportion of cost-of-living adjustment clauses in major collective bargain-
ing agreements was much higher in the 1970s and first half of the 1980s than the preceding
and subsequent periods. He finds that this pattern can be explained by a sizable increase of
inflation uncertainty in the 1970s, measured by the mean squared forecast error of inflation
surveys. Hofmann et al. (2012) estimate the evolution of US wage dynamics over time, and
find a degree of wage indexation to past inflation of 0.91 during the "Great Inflation"
of the 1970s, compared to 0.30 and 0.17 before and after this period.¹³ Hofmann et al.
(2012) argue that this evolution can be explained by a shift in the monetary policy reaction
function of the Federal Reserve. More specifically, a weakly inflation stabilizing conduct of
monetary policy in the 1970s resulted in high and volatile inflation, which encouraged the
use of indexation clauses in wage contracts as a protection against inflation uncertainty.
Conversely, the credible establishment of price stability after the disinflation of the early
1980s reduced the need for protection against unforeseen inflation, thus mitigating wage
indexation.

Both theory and empirical evidence thus suggest that the assumption of a constant
degree of wage indexation in a wage Phillips curve is questionable. However, although the
degree of wage indexation in the US was definitely higher during the "Great Inflation"
compared to earlier and later periods, it is not clear whether this was indeed the conse-
quence of monetary policy. For example, the 1970s were also characterized by very volatile
supply shocks, whilst changes in labor market institutions may also have played a role. It
is also not clear whether a link between indexation and the conduct of monetary policy
can be established in other countries. In the rest of this paper, we formally examine the
influence of the monetary policy regime on wage indexation within our panel dataset of
24 OECD countries.

Monetary policy regimes

We identify the monetary policy regime by the presence of an explicit quantitative mone-
tary target. It is commonly accepted that a policy regime that clearly commits to a nom-

¹³Hofmann et al. (2012) first estimate a time-varying parameters Bayesian structural vector autoregres-
sive (TVP-BVAR) model, and document considerable time variation in the impulse responses of wages
and prices to aggregate supply and demand shocks. In a second step, the parameters of a standard DSGE
model containing a wage Phillips curve are estimated for respectively 1960Q1, 1974Q1 and 2000Q1, by
matching the impulse responses from the TVP-BVAR for these periods with the impulse responses of the
DSGE model using a Bayesian impulse response matching procedure.
inal anchor can help promote price stability and stabilize inflation expectations (Mishkin, 2007). A quantitative monetary target should help to lower inflation uncertainty. Accordingly, also the degree of wage indexation is expected to be lower in a monetary policy regime with an explicit nominal anchor. The advantage of defining a policy regime by the presence of a quantitative target is that it can easily be measured and verified in an objective and mechanistic way (Fatás et al., 2007).

There are different types of nominal anchors, and in this work, we consider three different monetary target strategies: the inflation rate, the exchange rate and the money supply. These three frameworks have distinct characteristics. Inflation targeting provides a rational way to control inflation because policy decisions are based on conditional medium-run inflation forecasts. Its high degree of transparency and accountability further allows for close monitoring by the private sector (Svensson, 1999). A fixed exchange rate is easy to communicate to the private sector, but it cannot guarantee a strong policy commitment: monetary policy cannot react to domestic shocks independently from the anchor country, which makes the fixed exchange rate difficult to maintain under international capital mobility and leaves the country vulnerable to speculative attacks. Money growth targeting also offers immediate signals to the general public, but this strategy should be seen primarily as a way to communicate a commitment to low and stable inflation in the long run (Issing, 1996). A money aggregate is a less efficient predictor of future inflation in the short to medium term due to the unstable relationship between inflation and money aggregates. Broad monetary targets are also not under the direct control of a central bank.

The literature ascribes some different macroeconomic effects to the monetary regime types.\textsuperscript{14} Empirical evidence has found that inflation targeting has led to lower and less volatile inflation in developing economies, but these effects have not been noted for industrial countries (Walsh, 2009). Inflation targeting further seems to have anchored inflation expectations (Walsh, 2009) and lowered inflation persistence, which strengthens the nominal anchor (Mishkin and Schmidt-Hebbel, 2007).\textsuperscript{15} Although fixed exchange rate regimes have been linked to greater output volatility and lower and more stable inflation (Ghosh, Gulde, Ostry and Wolf, 1997), no significant effects have been reported for industrial countries (Levy-Yeyati and Sturzenegger, 2001). Monetary target regimes have been found to

\textsuperscript{14}See Fatás et al. (2007) and the references therein for further reading.

\textsuperscript{15}Gürkaynak, Levin and Swanson (2010); Gürkaynak, Levin, Marder and Swanson (2007) and Levin et al. (2004) in particular provide evidence for better anchoring of inflation expectations under inflation targeting.
keep inflation under control in the longer run by means of a flexible approach towards the
target rule and an active and elaborate communication of the monetary policy strategy
to the public (Mishkin, 1999). Finally, several countries have also achieved low and stable
inflation with hybrid targets (e.g. the EMU countries under the European Central Bank
and, formerly, Germany) and with implicit targets (most notably, the US until 2012).

Figure 4.1 summarizes the evolution of the monetary regimes for the countries in our
sample. The overall message is consistent with the shifts in policy regimes documented
in the literature (Fatás et al., 2007). Exchange rate targeting was the dominant regime
during the 1960s and early 1970s, but in the mid-1970s some countries shifted to monetary
targeting.\footnote{Exchange rate frameworks characterized as a managed or free float are not considered as an exchange rate targeting regime.} For a couple of decades, there are only minor changes. Then, in the early
1990s, inflation targeting enters the picture. This framework continuously gains ground
and becomes the dominant policy framework at the end of the sample. An important event
is the emergence of the European Central Bank for the EMU members in 1999. Our data
classifies EMU members as having both an inflation and a monetary target since 1999,
which is why the series jump at this point. Note also that the total number of regimes is always larger than the total number of countries, which indicates the combination of different targets, and that it rises over time while the share of regimes without an explicit target became negligible in the last decade.

Table 4.2 shows in detail how often one policy regime was abandoned for another, a move to which we refer as a regime switch. There are 55 of these regime switches in total, or more than 2 on average per country.\textsuperscript{17} It appears that countries often switched from exchange rate targeting to a regime without a target before finally settling on monetary or inflation targeting. If we look at the number of instances where there is a change in the policy regime without abandoning the regime already in place (including cases where, for instance, a monetary targeter adds an inflation target as second objective), we obtain a total of 70 regime changes. All countries, except Denmark, have changed their policy objective at some point in time; some countries have even done so for 6 times. We conclude that there is quite some variation in the data in terms of policy regime shifts, which demonstrates the advantage of pooling the data.

### 3.2 Econometric model specification

We formally explore the role of the monetary policy regime for wage indexation by extending the benchmark wage Phillips curve of section 2. Specifically, we allow the degree of wage indexation to vary according to the monetary policy regime, as well as some labor market characteristics, by estimating a multiplicative interaction wage Phillips curve

\textsuperscript{17}Of these 55 cases, only 6 are "double counted" cases where a country with one target switched to a regime with 2 other targets, and vice versa.
model that has the following form:

\[
\pi_{i,t}^w = \alpha_i + \alpha' D_{i,t} + \gamma' \pi_{i,t-1}^{p} + \gamma' D_{i,t} \pi_{i,t-1}^p + \psi_0 u_{i,t} + \psi_1 u_{i,t-1} + \eta_{i,t},
\]

(4.5)

where \( D_{i,t} \equiv (\text{REGIME}_{i,t}, \text{UD}_{i,t}, \text{COORD}_{i,t})' \)

\( D_{i,t} \) is a \( k \times 1 \) vector of variables that are interacted with past inflation and the constant \( \alpha_i \), whereas \( \alpha' \) and \( \gamma' \) are \( k \times 1 \) vectors with the corresponding interaction coefficients.\(^{18}\)

The policy regime interaction dummies (\text{REGIME}) indicate the presence of an explicit quantitative monetary target (\text{TARGET}), or respectively an inflation target (\text{IT}), a money growth target (\text{MON}) or an exchange rate target (\text{ER}).

We control for possible effects of changes in the wage bargaining process on indexation by including the coordination level of wage bargaining (\text{COORD}) and the union density rate (\text{UD}) in the estimations. Cecchetti (1987), for instance, documents that policy interventions in the bargaining process in the US during the 1960s and early 1970s altered both the effective degree of indexation and the frequency of union wage changes. Recent empirical evidence of Gnocchi and Pappa (2013) shows that wage bargaining reforms that reduced the centralization of wage bargaining have led to changes in wage dynamics. Messina and Sanz-de-Galdeano (2014) relate the different evolution of wage indexation in Uruguay and Brazil to the dynamics in the centralization level of wage bargaining and changes in the union coverage. The authors find that a decline in union coverage and a more decentralized wage bargaining reduces wage indexation. Carrillo et al. (2014) further show that the economy’s equilibrium degree of wage indexation can differ depending on whether the labor market coordination is centralized or decentralized.

We use the coordination level of wage bargaining to capture the degree to which major institutional players’ decisions extend to lower-level institutions and the percentage of workers that are affiliated with a labor union to control for the power of labor unions in wage bargaining negotiations. The coordination variable ranges from 1 (decentralized) to 5 (highly centralized), while the union density rate is expressed as a percentage. Figure 4.2 shows the evolution over time in the coordination measure for the countries in our sample. The proxy varies strongly across countries and time, and there is also a difference in the extent of time variation across countries (compare, e.g., Norway to the US). The same message holds for the union density rate, depicted in figure 4.3, although most countries seem to experience a downward or stable trend over time.

\(^{18}\)Interacting with the intercept prevents the estimation of spurious interaction effects (Brambor, Clark and Golder, 2006).
Figure 4.2 Degree of wage bargaining coordination (index: 1-5)

Figure 4.3 Union density rate
3.3 Results

We now discuss the estimation results of the wage Phillips curve with interaction variables, for which we only report FE estimates. Although one can motivate the CCEP estimator for the benchmark model, it is not suitable for the interaction effects model. The reason for this limitation is explained in table 4.3, which gives an overview of the mean and median correlation of the country-specific regime and labor market indicators with their respective CSA for the entire group of countries. The correlations are quite high, attaining 90% and more for the inflation target dummy, which indicates that changes in the interaction variables are synchronized in time across countries. This result makes controlling for their CSAs unattractive, as they will interfere with our goal of measuring the influence of the policy regime on $\gamma$. In addition, we found that including the CSAs of the other (non-interacted) variables does not lead to further important reductions of the extent of cross-sectional residual correlation. We therefore decided not to apply the CCE estimators for the estimation of the interaction model.

Table 4.3 Cross-country correlations

<table>
<thead>
<tr>
<th>Correlation interaction variables with their CSA</th>
<th>IT</th>
<th>ER</th>
<th>MON</th>
<th>COORD</th>
<th>UD</th>
</tr>
</thead>
<tbody>
<tr>
<td>mean</td>
<td>0.88</td>
<td>0.40</td>
<td>0.78</td>
<td>0.29</td>
<td>0.61</td>
</tr>
<tr>
<td>median</td>
<td>0.92</td>
<td>0.51</td>
<td>0.81</td>
<td>0.32</td>
<td>0.78</td>
</tr>
</tbody>
</table>

Table 4.4 reports the estimates of equation (4.5) with the monetary regime interaction variables. At the bottom of the table, we report the total indexation coefficient under both values of the binary monetary regime dummy variables.\(^{19}\) The first column indicates that monetary policy regimes with a quantitative monetary target, irrespective of the type of target, have a degree of wage indexation to past inflation which is on average not significantly lower than regimes without a quantitative target. Note however that the number of countries in our sample with an implicit target or an opaque monetary regime has become increasingly smaller over time and is especially limited since the 1990s (see figure 4.1). The results could by consequence be sensitive to the magnitude of the comparison group of countries with no formal and explicit monetary target.

A closer look at the effect of an inflation, money growth or exchange rate target, pro-

\(^{19}\)The total indexation attains $\frac{\partial \pi_{it}^w}{\partial \pi_{it-1}^p} = \gamma + \frac{5}{D_{it}}$. Standard errors are computed with the Delta method.
vided in column 2, shows that the effects differ. Specifically, the degree of wage indexation is on average 0.60 lower in policy regimes that have an explicit inflation target. With the labor market variables set to their sample means, wage indexation to past inflation is even statistically insignificant for countries with a quantitative inflation target. In contrast, a money growth or exchange rate target has no substantial effect on the indexation parameter.\textsuperscript{20} The level of coordination and the union density rate exert no significant influence on $\gamma$.\textsuperscript{21}

The result that an inflation target lowers wage indexation is consistent with Benati (2008), who finds that price indexation vanishes in countries with a stable nominal anchor. Inflation targets, unlike exchange rate and money targets, directly represent the ultimate goal of monetary policy and have been associated with strong nominal anchors (supra, p.124). The stability of a regime can also drive different effects, as more durable regimes produce better inflation outcomes (Rose and Mihov, 2008). Money growth and exchange rate regimes have been relatively short-lived in general. By contrast, no central bank has yet abandoned the inflation targeting framework.

\textsuperscript{20}The substantial reduction of indexation under inflation targeting is robust to using the first or fourth lag of the regime dummies. The effect is also robust to using the wage gap specification. These results are available from the authors on request.

\textsuperscript{21}We also tested for a possible nonlinear relationship of the extent of $\text{COORD}$ and $\text{UD}$ on wage indexation in the spirit of Calmfors and Drifill (1988), but the quadratic effects were found to be statistically insignificant.
### Table 4.4 Results interaction model

**Regressand:** wage inflation ($\pi_{it}^w$)

<table>
<thead>
<tr>
<th></th>
<th>(1)</th>
<th>(2)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Unemployment rate ($u_{it}$)</td>
<td>-0.370***</td>
<td>-0.334***</td>
</tr>
<tr>
<td></td>
<td>(0.089)</td>
<td>(0.092)</td>
</tr>
<tr>
<td>Lagged unemployment ($u_{it-1}$)</td>
<td>0.247***</td>
<td>0.214**</td>
</tr>
<tr>
<td></td>
<td>(0.087)</td>
<td>(0.091)</td>
</tr>
<tr>
<td>Lagged price inflation ($\pi_{it-1}^p$)</td>
<td>0.979***</td>
<td>0.961***</td>
</tr>
<tr>
<td></td>
<td>(0.138)</td>
<td>(0.109)</td>
</tr>
</tbody>
</table>

**Interactions:**

<table>
<thead>
<tr>
<th></th>
<th>(1)</th>
<th>(2)</th>
</tr>
</thead>
<tbody>
<tr>
<td>COORD</td>
<td>0.015</td>
<td>0.010</td>
</tr>
<tr>
<td></td>
<td>(0.027)</td>
<td>(0.021)</td>
</tr>
<tr>
<td>UD</td>
<td>-0.316</td>
<td>-0.350</td>
</tr>
<tr>
<td></td>
<td>(0.253)</td>
<td>(0.230)</td>
</tr>
<tr>
<td>TARGET</td>
<td>-0.157</td>
<td>-</td>
</tr>
<tr>
<td></td>
<td>(0.114)</td>
<td></td>
</tr>
<tr>
<td>IT</td>
<td>-</td>
<td>-0.598***</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(0.224)</td>
</tr>
<tr>
<td>MON</td>
<td>-</td>
<td>-0.031</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(0.083)</td>
</tr>
<tr>
<td>ER</td>
<td>-</td>
<td>-0.126</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(0.096)</td>
</tr>
</tbody>
</table>

Total indexation effect for (binary) MP regime dummies

<table>
<thead>
<tr>
<th></th>
<th>(1)</th>
<th>(2)</th>
</tr>
</thead>
<tbody>
<tr>
<td>TARGET =0/1 [coord&amp;ud at means]</td>
<td>0.897***/0.740***</td>
<td>-</td>
</tr>
<tr>
<td>IT=0/1 [coord&amp;ud at means]</td>
<td>-</td>
<td>0.852***/0.254</td>
</tr>
<tr>
<td>MON=0/1 [coord&amp;ud at means]</td>
<td>-</td>
<td>0.852***/0.820***</td>
</tr>
<tr>
<td>ER=0/1 [coord&amp;ud at means]</td>
<td>-</td>
<td>0.852***/0.725***</td>
</tr>
<tr>
<td>CD test (average correlation)</td>
<td>0.078</td>
<td>0.052</td>
</tr>
<tr>
<td>CD test (statistic and p-value)</td>
<td>13.14</td>
<td>9.38</td>
</tr>
</tbody>
</table>

Note: *, **, *** denote significance at 10, 5, and 1% levels, respectively.

Robust standard errors in brackets. Sample: N=24, $\bar{T}=137$, min $T=43$, max $T=207$

$\pi_{it-1}^p$ coefficient shows level of backward-looking indexation when all interactions $D_{it} = 0.$
4 Conclusions

We have examined the standard assumption in New Keynesian DSGE models that wage indexation to past price inflation is invariant to policy regimes. In particular, we have estimated the reduced form empirical wage Phillips curve of Galí (2011) with a panel model for 24 advanced economies, and allowed the degree of backward-looking wage indexation to vary according to the monetary policy regime while controlling for the evolution of labor market institutions.

We find that wage indexation to past inflation varies across monetary policy regimes. Specifically, policy regimes that have an explicit quantitative inflation target are characterized by a lower degree of wage indexation. Discerning between three different types of explicit targets - an inflation, money or exchange rate target - makes it clear that the extent of wage indexation is significantly different (lower) in countries that have an inflation target whereas the effects of money and exchange rate targets are not significantly different from a regime without any formal quantitative target. These differences could be due to varying strengths of the nominal anchor under the different frameworks, as inflation targeting has been found to establish better anchored inflation expectations (Mishkin and Schmidt-Hebbel, 2007; Walsh, 2009), which, in turn, could strengthen the nominal anchor.

Overall, our results question the structural nature of hard-wiring a fixed degree of wage indexation in standard DSGE models. Our work corroborates and extends the finding of Hofmann et al. (2012) of substantial time variation in the degree of wage indexation for the US. It further shows that the monetary policy dependence of price indexation found by Benati (2008) can be extended to wage indexation. From a policy standpoint, our findings suggest that counterfactual policy simulations and the analysis of optimal monetary policy based on modern macroeconomic models are potentially misleading.

A caveat to our analysis is that we focus on changes in wage indexation and do not measure changes in contract duration, a parameter that could also have been affected by policy changes (Cecchetti, 1987). We believe that our work provides empirical support for research that aims to endogenize the extent of wage indexation in structural models. Some steps have already been taken by Wieland (2009); Carrillo et al. (2014), but further research in this direction is warranted.
Appendix A: data sources and construction of variables

**Hourly earnings:** average total earnings in manufacturing paid per employee per hour, index 2010=100, seasonally adjusted [OECD, Main Economic Indicators database, quarterly frequency].

Note: The earnings series constitutes wage rates plus overtime payments, bonuses and gratuities regularly and irregularly paid, remuneration for time not worked, and payments in kind. Not included are employer contributions to social security or insurance schemes and unfunded employee social benefits paid by employers. Adding these components to the earnings series delivers compensation rates.

**Unemployment rate:** the unemployment rate is defined as the ratio of the number of unemployed workers to the working population [OECD, Main Economic Outlook No.93, quarterly frequency].

Note: The German series prior to 1992 were extended backwards based on the growth rates of unemployment rate as percentage of civilian labor force for West-Germany (source: Bundesbank, series BBK01.USCY01).

**Prices:** consumer price index, all items, index 2010=100, seasonally adjusted by Census X12 [OECD, Main Economic Indicators database, quarterly frequency].

**Monetary policy quantitative target dummies:** 0-1 dummy variables, which indicate if a country has a formal inflation, exchange rate or monetary target in the respective time period. If an exchange rate regime is classified as having a managed or freely floating exchange rate, we consider it to have no formal target [quarterly frequency].

Our main sources for the classification of quantitative monetary targets are Fatás et al. (2007) and Houben (2000). An additional source for inflation targeting regimes is Rose (2007); for exchange rate and monetary targeting regimes, it is Borio and White (2003).

---

22We initially started from a data sample of 29 OECD countries over 1960Q1-2013Q2. Due to data availability of the labor market indicators, the sample was limited to 2011Q4 and 3 countries were dropped (Iceland, Israel and Mexico). We further eliminated Japan and Luxembourg from the sample as these countries are characterized by severe outliers concerning the coefficients on the unemployment rate variables.
Note: Conflicting dates were examined and remaining gaps were filled based on central banks’ websites, individual central bank reports and, where necessary, additional publications.

**Coordination:** indicator of degree of wage coordination ranging from 1 (fragmented wage bargaining, confined largely to individual firms or plants) to 5 (centralized bargaining by peak association(s), with or without government involvement) [ICTWSS database version 4.0 from Visser (2013) (Amsterdam Institute for Advanced Labour Studies), annual frequency, 1960-2011].

**Trade union density:** the trade union density rate is defined as the ratio of the number of wage and salary earners that are trade union members to the total number of wage and salary earners [OECD, Labour statistics, annual frequency, 1960-2011].

**Trend inflation:** for each country, we estimate trend inflation with the AR-Trend-bound model of Chan et al. (2013) (see appendix B). The model delivers estimates of a central bank’s time-varying inflation target, which is restricted to lie within bounds, based on diffuse uniform priors.

**Overview number of observations per country:**

<table>
<thead>
<tr>
<th>country</th>
<th>time span</th>
<th>#</th>
<th>country</th>
<th>time span</th>
<th>#</th>
</tr>
</thead>
<tbody>
<tr>
<td>Australia</td>
<td>1984Q2-2011Q4</td>
<td>111</td>
<td>Italy</td>
<td>1960Q2-2011Q4</td>
<td>207</td>
</tr>
<tr>
<td>Austria</td>
<td>1969Q2-2011Q4</td>
<td>171</td>
<td>Korea, South</td>
<td>1992Q2-2011Q4</td>
<td>79</td>
</tr>
<tr>
<td>Belgium</td>
<td>1960Q2-2011Q4</td>
<td>207</td>
<td>Netherlands</td>
<td>1970Q2-2011Q4</td>
<td>167</td>
</tr>
<tr>
<td>Canada</td>
<td>1960Q2-2011Q4</td>
<td>207</td>
<td>New Zealand</td>
<td>1989Q2-2011Q4</td>
<td>91</td>
</tr>
<tr>
<td>Czech Republic</td>
<td>1993Q2-2011Q4</td>
<td>75</td>
<td>Norway</td>
<td>1960Q2-2011Q4</td>
<td>207</td>
</tr>
<tr>
<td>Denmark</td>
<td>1971Q2-2011Q4</td>
<td>163</td>
<td>Poland</td>
<td>1995Q2-2011Q4</td>
<td>67</td>
</tr>
<tr>
<td>Estonia</td>
<td>2000Q2-2011Q4</td>
<td>47</td>
<td>Portugal</td>
<td>2000Q2-2011Q4</td>
<td>47</td>
</tr>
<tr>
<td>Finland</td>
<td>1960Q2-2011Q4</td>
<td>207</td>
<td>Slovak Republic</td>
<td>1993Q3-2011Q4</td>
<td>74</td>
</tr>
<tr>
<td>France</td>
<td>1960Q2-2011Q4</td>
<td>207</td>
<td>Spain</td>
<td>1981Q2-2011Q4</td>
<td>123</td>
</tr>
<tr>
<td>Germany</td>
<td>1962Q2-2011Q4</td>
<td>199</td>
<td>Sweden</td>
<td>1971Q2-2011Q4</td>
<td>163</td>
</tr>
<tr>
<td>Hungary</td>
<td>1995Q2-2011Q4</td>
<td>67</td>
<td>United Kingdom</td>
<td>1963Q2-2011Q4</td>
<td>195</td>
</tr>
<tr>
<td>Ireland</td>
<td>1990Q2-2011Q4</td>
<td>87</td>
<td>United States</td>
<td>1960Q2-2011Q4</td>
<td>207</td>
</tr>
</tbody>
</table>

Table 4.5 Overview number of observations per country
Appendix B: AR-Trend-bound model

This section provides additional details on the estimation of the AR-Trend-bound model (see section 2.1). Chan et al. (2013) propose a new model of trend inflation which restricts the central bank’s inflation target to vary between bounds. They argue that the commonly used random walk specification for the inflation target is counterintuitive because it implies that long-run expectations can grow in an unbounded fashion. Furthermore, they show that bounding trend inflation results in good long-term forecasting properties for inflation.

Model specification The AR-trend-bound model specifies that

\[
\begin{align*}
\pi_t^p - \pi_t^s &= \rho_t \left( \pi_{t-1}^p - \pi_{t-1}^s \right) + \epsilon_t \exp \left( \frac{h_t}{2} \right) \\
\pi_t^s &= \pi_{t-1}^s + \epsilon_t^s \\
h_t &= h_{t-1} + \epsilon_t^h \\
\rho_t &= \rho_{t-1} + \epsilon_t^\rho,
\end{align*}
\]

where \( \epsilon_t \sim N(0,1) \) and \( \epsilon_t^h \sim N(0, \sigma_h^2) \). Although trend inflation follows a random walk, the model restricts it to lie within bounds: \( a \leq \pi_t^s \leq b \ \forall t \). The \( \rho_t \) parameter governs the persistence of the inflation gap, i.e. the difference between inflation and the target, and it is restricted as \( a_\rho \leq \rho_t \leq b_\rho \ \forall t \).23 Given these restrictions, the innovations to the state equations of trend inflation and the persistence parameter are assumed to follow a truncated normal distribution:

\[
\begin{align*}
\epsilon_t^s &\sim TN(a - \pi_{t-1}^s, b - \pi_{t-1}^s; 0, \sigma_s^2) \\
\epsilon_t^\rho &\sim TN(a_\rho - \rho_{t-1}, b_\rho - \rho_{t-1}; 0, \sigma_\rho^2)
\end{align*}
\]

The Priors One can fix the bounds \((a, b)\) in advance or estimate them. We have set relatively diffuse uniform priors for the lower and upper bounds as \( a \sim U(a_\pi, \bar{a}) \) and \( b \sim U(b_\pi, \bar{b}) \).

\[23\]We have also experimented with the UC-SV model of Stock and Watson (2007), both with constant and stochastic volatility processes for the shocks in the law of motion of trend inflation. The problem is that the UC-SV model assumes that the inflation gap series to be independently distributed error terms (with stochastic variance). This implies a highly volatile inflation trend and causes past inflation and trend inflation to be highly collinear.
for each country (see table 4.6). Keep in mind that these bounds apply to quarter-on-quarter non-annualized growth rates of the consumer price index and that some countries experienced significantly high inflation rates.

Chan et al. (2013) initialize the state equations with

\[
\begin{align*}
\pi_1^* &\sim TN(a, b; \pi_0^*, \omega_{\pi^*}^2) \\
\rho_1 &\sim TN(a_\rho, b_\rho; \rho_0, \omega_\rho^2) \\
h_1 &\sim N(h_0, \omega_h^2),
\end{align*}
\]

with \(\pi_0^* = h_0 = \rho_0 = 0\), \(\omega_{\pi^*}^2 = \omega_h^2 = 5\), and \(\omega_\rho^2 = 1\) in order to ensure proper yet relatively uninformative prior distributions for the initial states. They set \(a_\rho\) and \(b_\rho\) such that \(0 < \rho_t < 1\ \forall t\). Since our analysis contains many countries, we opted to use diffuse priors that put even more weight on the sample information and can be applied for all our countries. Relative to their setup, we make initial conditions more diffuse by setting \(\omega_{\pi^*}^2 = \omega_h^2 = 25\) and set \(\pi_0^*\) equal to the mean of the first four observations of the inflation series. Furthermore, we only restrict the inflation gap to be stationary: \(-1 < \rho_t < 1\ \forall t\).

Chan et al. (2013) assume the following inverse gamma priors for the error variances:

\[
\begin{align*}
\sigma_{\pi^*}^2 &\sim IG(\nu_{\pi^*}, S_{\pi^*}) \\
\sigma_\rho^2 &\sim IG(\nu_\rho, S_\rho) \\
\sigma_h^2 &\sim IG(\nu_h, S_h),
\end{align*}
\]

with small (uninformative) degrees of freedom parameters \(\nu_{\pi^*} = \nu_\rho = \nu_h = 10\) and \(S_{\pi^*} = .18\), \(S_\rho = .009\), and \(S_h = .45\). For our analysis, we follow Cogley and Sargent (2005) in placing the most weight on the sample information by using an IG prior with a single degree of freedom for all three variances: \(IG(\frac{1}{2}, \frac{0.01^2}{2})\).

Our estimation uses Bayesian MCMC methods, as detailed in Chan et al. (2013), and we drew 55,000 samples and discarded the first 5,000 as burn-in. The results are available upon request.
### Table 4.6 Uniform prior settings for the trend-inflation bounds

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<th>$\bar{b}$</th>
<th>Country</th>
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<th>$\bar{a}$</th>
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Bibliography


