Essays in Macroeconomics

by

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Abstract (English)

This thesis consists of three chapters. The first chapter examines empirically the relationship between foreign aid and economic growth in the Least Developed Countries. Instrumental variables techniques are used to estimate the effect that economic growth has on foreign aid and to adjust for the reverse causal effect that growth has on aid when estimating the effect that aid has on growth. The second chapter examines the effects that fiscal expansions have on the unemployment rate. The chapter presents SVAR evidence for ten OECD countries and builds a DSGE model with a labor force participation choice and workers' heterogeneity to explain the empirical findings. The third chapter examines the effects that economic growth has on the support for extreme political platforms. The chapter provides a theoretical model in favor of growth effects (as opposed to level effects) on the support for extreme political parties, and investigates empirically the relationship between growth and extremist votes for 16 OECD countries.

Abstract (Spanish)

Esta tesis consiste en tres capítulos. El primer capítulo examina empíricamente la relación entre la ayuda exterior y crecimiento económico en los países menos adelantados. Técnicas de variables instrumentales se utilizan para estimar el efecto que el crecimiento económico tiene sobre la ayuda exterior y para ajustar el efecto de causalidad inversa que el crecimiento tiene en la ayuda al estimar el efecto que la ayuda tiene sobre el crecimiento. El segundo capítulo analiza los efectos que las expansiones fiscales tienen sobre la tasa de desempleo. El capítulo presenta pruebas SVAR para diez países de la OCDE y construye un modelo DSGE con una participación en la fuerza de trabajo y heterogeneidad de los trabajadores para explicar los resultados empíricos. El tercer capítulo analiza los efectos que el crecimiento económico tiene en el apoyo a las plataformas políticas extremas. El capítulo ofrece un modelo teórico a favor de los efectos del crecimiento (en contraposición a los efectos de nivel) con el apoyo de partidos políticos de extrema, e investiga empíricamente la relación entre el crecimiento de votos y extremistas para 16 países de la OCDE.

Preface

This thesis consists of three chapters. The first chapter examines empirically the relationship between foreign aid and economic growth in the Least Developed Countries. The chapter shows that foreign aid has a significant positive average effect on real per capita GPD growth if, and only if, the quantitatively large negative reverse causal effect of per capita GDP growth on foreign aid is adjusted for in the growth regression. Instrumental variables estimates yield that a 1 percentage point increase in GDP per capita growth decreased foreign aid by over 4 percent. Adjusting for this quantitatively large, negative reverse causal effect of economic growth on foreign aid yields that a 1 percent increase in foreign aid increased real per capita GDP growth by over 0.1 percentage points. The chapter shows that the obtained instrumental variables estimates are consistent with a calibrated version of a Solow-Swan growth model where part of foreign aid is used for investment. The chapter also examines cross-country parameter heterogeneity and distinguishes between short-run and long-run effects of foreign aid on economic growth.

The second chapter examines the effects that fiscal expansions have on the unemployment rate. Structural VARs indicate that, for many OECD countries the unemployment rate significantly increases following increases in government expenditures under a variety of specifications and identification schemes. Fiscal expansions also tend to increase employment, participation rates and real wages. Existing models have difficulties in generating such responses. The chapter shows that the empirical regularities can be reproduced with two additions into a standard New Keynesian model with matching frictions: (a) a labor force participation choice and (b) workers' heterogeneity. The chapter is joint work with Evi Pappa.

The third chapter examines the effects that economic growth has on the support for extreme political platforms. The chapter provides a theoretical argument in favor of growth effects (as opposed to level effects) on the support for extreme political parties and empirically investigates the relationship between growth and extremist votes. Lower growth rates benefit right-wing and nationalist parties, but do not have a robust positive effect on the support for communist parties. The empirical estimates indicate that extreme political platforms are unlikely to gain majorities in OECD countries, unless there is an extreme drop in the GDP per capita growth rate. The chapter is joint work with Hans Grüner.

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Chapter 1: On the Simultaneity Problem in the Aid and Growth Debate

1 Introduction

Does foreign aid have a positive, causal effect on economic growth? I show that the answer to this important policy question is yes if, and only if, one takes into account that economic growth itself has a quantitatively large, negative within-country effect on foreign aid. The aid effectiveness literature is well aware of this endogeneity problem. However, one of the main problems that this literature continues to struggle with, is finding a plausible time-varying instrumental variable for foreign aid (Temple, 2010). Moreover, despite standard macroeconomic theory predicting a positive effect of foreign aid on economic growth if part of the foreign aid is used for investment, the consensus in the aid effectiveness literature is that foreign aid does not have a significant positive average effect on economic growth.[1] I show that indeed one may arrive at this conclusion if the negative reverse causal effect of per capita GDP growth on foreign aid is not accounted for in the growth regression. Once the negative reverse causal effect of economic growth on foreign aid is accounted for, estimates of the effect of foreign aid on economic growth are positive, statistically significant, and economically meaningful.

My estimation strategy to identify the causal effect of foreign aid on economic growth is based on a two-step procedure. The two-step procedure is closely related to the approach taken in the empirical macro literature to identify the causal effects of fiscal policy (see, in particular, Blanchard and Perotti, 2002). In the first step, I estimate the response of foreign aid to economic growth, using rainfall and international commodity price shocks as instrumental variables to generate exogenous variation in real per capita GDP growth for a panel of 47 LDCs during the period 1960-2000. In the second step, after the causal response of foreign aid to real per capita GDP growth is quantified by the instrumental variables estimates, I use the residual variation in foreign aid that is not driven by GDP per capita growth as an instrument to estimate by two-stage least squares the effect that foreign aid has on per capita GDP growth. This two-step estimation strategy enables to: (i) obtain an understanding of how foreign aid responds to per capita GDP growth at the macroeconomic level (hence, providing useful information on the severity of the endogeneity bias if one fails to adequately deal in the growth regression with the endogenous response of foreign aid to economic growth); and (ii) compute an estimate of the effect that foreign aid has on economic growth that is adjusted for the reverse causal effect that growth has on aid.

[1]See for example Burnside and Dollar (2000), Hansen and Tarp (2001), Dalgaard et al. (2004), Easterly et al. (2004), Roodman (2007), Bourguignon and Sundberg (2007), or Rajan and Subramanian (2008), and the critical review of the literature by Temple (2010). Papers that have studied aid allocation criteria include, among many others, Trumbull and Wall (1994), Alesina and Dollar (2000), and Alesina and Weder (2002).

My first main finding is that increases in real per capita GDP growth of aid recipient countries are associated with a statistically significant and quantitatively large reduction in foreign aid. An instrumental variables estimate yields that a 1 percentage point increase in the real per capita GDP growth rate is associated with a significant decrease in foreign aid by over 4 percent. This result is consistent with the stylized cross-country fact that as countries grow richer they rely less on foreign aid. It is also consistent with donor countries acting as Good Samaritans: when the economy of the aid recipient country is booming the Good Samaritan reduces aid, while in times of severe economic difficulties aid is increased.

An important implication of this first main finding is that research on the effect of foreign aid on economic growth is complicated by a quantitatively large, negative reverse causal effect of economic growth on foreign aid. The large, negative causal effect of economic growth on foreign aid implies that the cards in empirical research on aid effectiveness are stacked against finding in the data a significant positive average effect of foreign aid on economic growth.

I show that once the negative reverse causal effect of per capita GDP growth on foreign aid is adjusted for in the growth regression, that foreign aid did indeed have a significant positive average effect on real per capita GDP growth. My panel fixed effects estimates yield that a 1 percent increase in foreign aid is associated with a significant within-country increase in GDP per capita growth of around 0.1 percentage points. These instrumental variables estimates are consistent with the quantitative predictions of a Solow-Swan growth model where a part of foreign aid finances domestic investment. I also show that estimating the effect that foreign aid has on GDP per capita growth without taking into account that there is a large, negative reverse causal effect of economic growth on foreign aid would lead to the (mistaken) conclusion that foreign aid has no significant positive average effect on real per capita GDP growth.

There are several reasons why the issue of the causal effect that foreign aid has on economic growth in the Least Developed Countries is important. First, foreign aid flows constitute a significant share of these countries' per capita income. On average the 1960-2000 share of net official development aid in GDP was about 3 percent. Second, most of the foreign aid flows are financed by tax payer money. Western governments are accountable to voters and most of these voters may not be receptive to much more information than whether on average foreign aid had a positive causal effect on economic growth. Third, the premier World Millenium Development goal is to end poverty and hunger in the world's poorest countries. If one believes that per capita GDP growth is associated with significant increases in the income per capita of the world's poorest, as is suggested for example by Dollar and Kraay (2002), and if one cares about reducing poverty and hunger, then clearly it is important to have kowledgement about the causal effect that foreign aid has on economic growth in the world's poorest countries.

The remainder of the paper is organized as follows. Section 2 explains the estimation strategy. Section 3 describes the data. Section 4 presents the main results. Section 5 presents further robustness checks. And Section 6 concludes.

2 Estimation Strategy

2.1 Estimating the Effect that Economic Growth has on Foreign Aid

Estimating the effect that real per capita GDP growth has on foreign aid requires an exogenous source of variation for real per capita GDP growth. To generate such variation, I use smooth variations in rainfall and international commodity price shocks as instrumental variables.[2] A key characteristic of the Least Developed Countries (LDCs) that makes this estimation strategy plausible is that these countries are highly dependent on the agricultural and commodity exporting sector.[3] Hence, variations in rainfall and international commodity prices can induce substantial variation in real per capita GDP growth vis-a-vis changes in agricultural productivity and the terms of trade. Because rainfall is random and the economic size of each LDC (as measured by the share in world commodity production) is extremely small (so that the country can be effectively treated as being a price taker on the international commodity market) the induced variations in per capita GDP growth will be exogenous to variations in foreign aid and economic growth.[4]

I estimate the effect that real per capita GDP growth has on foreign aid using two-stage least squares:

$$\Delta log(aid_{i,t}) = a_i + b_t + c\Delta log(y_{i,t}) + e_{i,t}, \tag{1}$$

where $\Delta log(aid_{i,t})$ is the log-change of foreign aid per capita and $\Delta log(y_{i,t})$ is the log-change of real per capita GDP.[5] a_i are country fixed effects that capture long-run (unobservable) differences across countries that jointly determine changes in foreign aid per capita and per capita GDP growth; b_t are year fixed effects that capture global business cycle effects and other global shocks that may be jointly driving foreign aid and per capita GDP growth of the LDCs.

[2] Several papers have documented the significant effect of rainfall and international commodity price shocks on economic growth in Sub-Saharan Africa. See for example Deaton (1999), Miguel et al. (2004), or Brückner and Ciccone (2010a,b). Sub-Saharan African countries constitute about two-thirds of the 49 countries that are classified by the United Nations as the Least Developed Countries (LDCs). The paper covers 47 of the 49 LDCs. The 2 LDCs that are not covered in the paper due to missing GDP data are East-Timor and Myanmar.

[3] See the Data Appendix for further details.

[4] Conditional of course on country and year fixed effects. See the equation below.

[5] This functional form follows Trumball and Wall (1994), who derive the panel fixed effects log-log specification based on a theoretical model where aid decisions of donors are motivated by the well-being of the aid recipient country. I use the log-change of foreign aid rather than the level of foreign aid because the Im, Pesaran, and Shin (2003) panel unit root test did not reject the null hypothesis that the level of foreign aid has a unit root. The test rejected however at the 1% level the null hypothesis that the first-difference of the foreign aid series has a unit root. Regarding cointegration, the panel cointegration tests developed by Westerlund (2007) did not reject the null hypothesis that there is no cointegration between the log-level of foreign aid and the log-level of GDP per capita. Thus, panel cointegration tests do not point to a significant exact common component between permanent shocks to the level of GDP per capita and permanent shocks to the level of foreign aid.

The excluded instruments ($Z_{i,t}$) in the two-stage least squares estimation of equation (1) are the log-changes of the international commodity price index, rainfall, and rainfall squared (see Section 3 for a detailed description of how these instruments are constructed). The exclusion restriction states that the instruments should only systematically affect the dependent variable (foreign aid per capita) through their effect on per capita GDP growth. The validity of rainfall and commodity price shocks as excluded instruments in equation (1) will be examined rigorously in Section 4.1.

2.2 Estimating the Effect that Foreign Aid has on Economic Growth

If per capita GDP growth has a significant effect on foreign aid (i.e. in equation (1) c $\neq 0$) then OLS estimation of the effect that foreign aid has on economic growth will be biased. Specifically, suppose that the effect of foreign aid on economic growth can be written as:

$$\Delta log(y_{i,t}) = h_i + i_t + k\Delta log(aid_{i,t}) + mZ_{i,t} + u_{i,t}, \tag{2}$$

then $cov(\Delta log(aid_{i,t}), u_{i,t}) \neq 0$, and OLS estimation of k will be upward biased if $c > 0$ and downward biased if $c < 0$.

This endogeneity bias, that is due to $c \neq 0$ in equation (1) can be evaded however by (i) constructing an adjusted foreign aid series where the response of foreign aid to per capita GDP growth is partialled out; i.e.:

$$\Delta log(aid_{i,t})^* = \Delta log(aid_{i,t}) - c\Delta log(y_{i,t}) \tag{3}$$

and (ii) using this endogeneity adjusted aid series as an instrument for the original aid series in equation (2). By construction, the IV estimator that uses the endogeneity adjusted aid series $\Delta log(aid_{i,t})^*$ as an instrument for $\Delta log(aid_{i,t})$ does not suffer from the simultaneity bias. Moreover, beyond taking care of the simultaneity bias, that is associated with the least squares estimation of equation (2), the IV estimator will provide a consistent estimate of the parameter k under the assumption (exclusion restriction) that the error in equation (1) is uncorrelated with the error in equation (2). If there are omitted variables that are part of both, equations (1) and (2) the zero-covariance assumption will be violated and the IV estimator will not solve the omitted variables problem. However, the IV estimator will still solve the simultaneity problem. The Technical Appendix provides a formal proof for why an IV estimator that uses the residual variation of foreign aid which is not driven by economic growth does not suffer from the simultaneity bias. The appendix also derives the omitted variables bias of the least squares and IV estimator which arises when the zero-covariance restriction is violated.

Note that the estimation strategy requires that the parameter c in equation (1) is estimated consistently. Because of the simultaneous nature of the two equations, OLS can not provide a consistent estimate of the parameter c in equation (1) if $k \neq 0$ in equation (2). Moreover, because measurement error is a real concern in national accounts statistics of developing countries (e.g. Heston, 1994; Deaton, 2005) the OLS estimate of the parameter c in equation (1) will likely be attenuated towards zero. Hence, the need for instrumental variables estimation of equation (1).

An issue arising with the estimation strategy in equation (2) is that the adjusted aid series $(aid_{i,t})^*$ is a generated regressor. Typically, the presence of a generated regressor leads to standard errors on the slope coefficients that are incorrect for purposes of hypothesis testing.[6] However, there is a special case where the standard error on the slope coefficient of a generated regressor is correct: namely, for testing the hypothesis that the slope coefficient is equal to zero (see, for example, Wooldridge, 2002, p. 141). In the aid literature the debate has focused on the question of whether the effect of foreign aid on economic growth is significantly different from zero. Hence, the special case where the standard error on the slope coefficient of a generated regressor is correct (i.e. the case for testing the hypothesis of a zero slope coefficient on foreign aid) is the relevant one for this paper's empirical analysis.

3 Data

Rainfall Data. I obtain data on annual rainfall for each of the 47 LDCs during the period 1960-2000 from the Climate Research Unit (CRU) and the Tyndall Centre for Climate Change Research (TYN) of the University of East Anglia. Specifically, I use the TYN CY 1.1 version that has been developed by Mitchell et al. (2003) and approved by the Intergovernmental Panel on Climate Change (IPCC). The CRU/TYN rainfall data come at a high resolution (0.5°x0.5° latitude-longitude grid) and each rainfall observation in a given grid is constructed by interpolation of rainfall observed by all stations operating in that grid. Rainfall data are then aggregated to the country level by assigning grids to the geographic borders of countries and weighting the observation in each grid by surface area, using the cosine of the latitude (see Mitchell et al., 2003 for more details).

[6] Consistency of the estimator is of course not affected by the use of a generated regressor.

International Commodity Price Shocks. The country-specific international commodity export price index $ComPI_{i,t}$ that captures shocks to the international prices of exported commodities is constructed as:

$$ComPI_{i,t} = \prod_{c \in C} ComPrice_{c,t}^{\theta_{c,i}} \quad (4)$$

where $ComPrice_{c,t}$ is the international price of commodity c in year t, and $\theta_{c,i}$ is the average (time-invariant) value of exports of commodity c in the GDP of country i. Annual international commodity price data are for the 1960-2000 period from UNCTAD Commodity Statistics, and data on the value of commodity exports are from the NBER-United Nations Trade Database.[1]

GDP and Foreign Aid Data. The real per capita GDP data are from the Penn World Tables (PWT), version 6.2 (Heston et al., 2006). Data on net official development aid are from the World Development Indicators (2009).[2]

4 Main Results

4.1 IV Estimates of the Effect of Economic Growth on Foreign Aid

Table 1 presents the baseline two-stage least squares estimates of the effect that real per capita GDP growth has on foreign aid. Column (1) shows the first-stage estimates that link international commodity price shocks and rainfall ($Z_{i,t}$) to real per capita GDP growth $\Delta log(y_{i,t})$. All three instruments are individually significant at least at the 1% level and yield a first-stage F-statistic of about 9.3. Increases in the international prices of exported commodities and improved

[1] The commodities included in the index are: aluminum, beef, coffee, cocoa, copper, cotton, gold, iron, maize, oil, rice, rubber, sugar, tea, tobacco, wheat, and wood. In case there were multiple prices listed for the same commodity a simple average of all the relevant prices is used.

[2] I use net official development aid, defined as grants and concessional loans net of repayments, because this measure captures best the actual transfers to countries (see for example Easterly, 2003, p. 29). I have chosen to focus on total official development aid, rather than more specific measures of aid, because if aid is fungible, as argued for instance in Devarajan and Swaroop (1998), then conceptually it makes little sense to distinguish between different kinds of foreign aid (see also Rajan and Subramanian, 2008).

rainfall conditions are associated with a significant increase in the real per capita GDP growth of the LDCs. The negative quadratic term on the rainfall variable captures that at some stage too much rainfall may be counterproductive for agricultural productivity and hence for GDP per capita growth.

That the instruments $Z_{i,t}$ have also a significant reduced-form effect on foreign aid $\Delta log(aid_{i,t})$ is shown in column (2). Increases in the international prices for exported commodities and improved rainfall conditions are associated with a significant decrease in foreign aid. Because the regression controls for year fixed effects, the reduced-form estimates are not driven by changes in economic conditions of OECD countries that may in turn systematically affect movements of international commodity prices. Moreover, the country fixed effects take into account that some LDCs are more dependent on the agricultural and commodity exporting sector than others, and that aid flows may be determined by deep historical factors, such as for example colonial ties to a specific European country.

For comparison purposes with the second-stage estimates that are presented in columns (4)-(8), column (3) shows the least squares estimates of the effect that real per capita GDP growth has on foreign aid. The least squares estimate is negative and statistically significant at the 5% level. However, if foreign aid has a significant effect on GDP per capita growth this point estimate can not be taken as reflecting the causal effect that per capita GDP growth has on foreign aid. In fact, if foreign aid has a significant positive effect on GDP per capita growth, the least squares estimate of the effect that GDP per capita growth has on foreign aid will be upward biased.

Column (4) therefore presents the two-stage least squares estimate that uses international commodity price shocks and rainfall as excluded instruments. The second-stage point estimate on real GDP per capita growth from the two-stage least squares regression is statistically significant at the 5% level and in absolute size much larger than the corresponding least squares estimate in column (3). The larger absolute size of the coefficient from the two-stage least squares regression could be due to a number of factors. First, if foreign aid has a positive effect on GDP per capita growth the least squares estimate will be upward biased. Second, measurement error in per capita GDP growth is a real issue for the LDCs (see for example, Heston, 1994; or Deaton, 2005). To the extent that this measurement error is classical it will attenuate the slope coefficient in the least squares regression towards zero but not the slope coefficient in the two-stage least squares regression. The Hausman test rejects that the least squares estimate is equal to the two-stage least squares estimate at the 10% level (p-value 0.067), thus pointing to a significant difference between the least squares and instrumental variables estimate.

Quantitatively, the two-stage least squares estimate in column (4) implies that a 1 percentage point increase in real per capita GDP growth is associated with an average reduction in foreign aid by over 4 percent. For this two-stage least squares estimate to reflect the causal effect that per capita GDP growth has on foreign aid, it is necessary that the instruments fulfill the exclusion restriction. That is, rainfall and international commodity price shocks should have no systematic effects on foreign aid other than through GDP per capita growth.

The p-value of the Hansen J-test on the overidentifying restrictions reported in column (4) is 0.82. Hence, the Hansen J-test does not reject that the instruments are uncorrelated with the second-stage error. To show also more intuitively that beyond per capita GDP growth there are no systematically large direct effects of international commodity price shocks and rainfall on foreign aid I report in columns (5)-(7) two-stage least squares estimates when instruments are added to the right-hand side of the second-stage equation. As can be seen, the size of the coefficient on the international commodity price index conditional on real per capita GDP growth (columns (5) and (6)) is less than one-third of the size of the coefficient that is obtained in the reduced-form regression (column (2)). Statistically, the coefficient is also not significant at any conventional confidence level. The coefficient on rainfall on the other hand flips sign and is also statistically insignificant. These regressions that directly estimate the effect that the instruments have on foreign aid conditional on per capita GDP growth therefore resonate the result of the Hansen J-test that did not reject the validity of rainfall and international commodity price shocks as instrumental variables for real per capita GDP growth in the aid equation. Column (8) shows that the second-stage (and first-stage) relationship continues to hold when excluding all those country-years where LDCs experienced a drought year, which could be associated with an atypical influx of foreign aid.[3]

An issue that has received substantial attention in particular in the aid literature is the robustness of results to outliers and the sample size.[4] To show that within the LDC sample results are robust to the selection of a specific sub-sample and the exclusion of observations that may be deemed as potential outliers Table 2 presents a variety of robustness checks. In column (1) only those country-year observations are used for the two-stage least squares estimation that produce a balanced panel for the 1960-2000 period. The point estimate on the second-stage coefficient is in this case -5.47 and is statistically significant at the 1% level. In column (2) the balanced sample is maintained, but all those observations are excluded which are deemed as outliers by the Hadi (1992) procedure.[5] Excluding these outliers barely changes the second-stage point estimate, but it does make the first-stage fit a bit more precise. In columns (3)-(5) the sample period is elevated to cover the 1970-2000 period only. The motivation for focusing on the 1970-2000 period is that some of the LDCs during the 60s were still under colonial influence of the European countries. Column (3) presents the unbalanced panel estimates for the 1970-2000 period, while column (4) uses only those 39 LDCs that yield a balanced panel for the

[3]Drought years are identified using the publicly available data on natural disasters that are provided by the Universite Catholic de Louvain's Emergency Disaster database (EM-DAT, 2009).

[4]See for example Easterly et al. (2004) or Roodman (2007).

[5]The Hadi (1992) procedure for detecting outliers has been popularly used in the aid literature. See for example Easterly et al. (2004) or Roodman (2007). The cut-off significance level chosen for the Hadi procedure is 5%.

1970-2000 period. Column (5) excludes further potential outliers based on the Hadi procedure. The main result is that per capita GDP growth continues to have a significant negative effect on foreign aid in all these regressions. Point estimates range between -4.19 and -6.07 and their 95% confidence intervals span the point estimate obtained in column (4) of Table 1.[6]

4.2 IV Estimates of the Effect that Foreign Aid has on Economic Growth

The results of the previous section showed that foreign aid is highly endogenous to the per capita GDP growth of the aid recipient countries. Specifically, the instrumental variables estimates yielded that foreign aid decreased substantially during times when per capita GDP growth of aid recipient countries increased. Hence, an OLS estimate of the effect that foreign aid has on per capita GDP growth will suffer from downward bias due to the reverse negative effect that per capita GDP growth has on foreign aid.

Panel A of Table 3 shows estimates of the effect of foreign aid on per capita GDP growth when adjusting for the large negative effect that GDP per capita growth has on foreign aid (for an explanation of how this is done see Section 2.2).[7] Panel B reports for comparison purposes the OLS estimates. All regressions continue to control for country and year fixed effects.

The main message of the estimates in Panel A of Table 3 is that the effect of foreign aid on real per capita GDP growth is positive and significantly different from zero at over 99% confidence when the negative reverse causal effect of GDP per capita growth on foreign aid is adjusted for. On the other hand, in Panel B of Table 3 the OLS estimates, that suffer from the negative reverse causal effect are either statistically insignificant or significantly negative. These results hold across a variety of different sub-sample specifications and are robust to the exclusion of observations that are deemed as potential outliers by the Hadi procedure. Specifically, the instrumental variables estimates in Panel A of Table 3 yield that a 1 percent increase in foreign aid is associated with a significant increase in real per capita GDP growth by around 0.1 to 0.2 percentage points.[8]

[6] An additional criterion that is important for instrumental variables estimation to yield consistent second-stage estimates is the first-stage relevance of the instruments. The first-stage F-statistic in Tables 1 and 2 is between 7.8 and 15.8. According to the tabulations in Stock and Yogo (2005), the maximal IV relative bias (maximal size distortion) is therefore less than 5% to 20% (15% to 25%). The p-values reported in square brackets below the 2SLS estimates in Tables 1 and 2 are based on the Anderson-Rubin test of statistical significance, and a key property of this test statistic is robustness to weak instruments (see for example Andrews and Stock, 2005). In Appendix Table 1 I show that using weak IV robust estimators yields second-stage estimates that are very similar, both quantitatively and statistically, to the two-stage least squares estimates reported in Tables 1 and 2.

[7] The adjustment is done using the corresponding point estimates of the effect that economic growth has on aid from Tables 1 and 2, thus matching the sample size in each column of Table 3.

[8] Appendix Table 2 shows that similar results are obtained when applying the instrumental variables strategy to the publicly available datasets of Burnside and Dollar (2000), Easterly et al. (2004), or Roodman (2007).

Table 4 shows that there continues to be a significant positive average effect of foreign aid on economic growth when controlling for within-country changes in political institutions. Changes in political institutions could have a direct and independent effect on foreign aid beyond economic growth if due to political reasons donors prefer to give foreign aid to more democratic countries.[9] For within-country changes in political institutions to be an omitted variable in the growth equation it would have to be the case however that at the annual level a change in political institutions has an immediate effect on economic growth. Panel A of Table 4 shows that this is not the case. The coefficient on the Polity2 score that captures political institutions is statistically insignificant and quantitatively small. The average marginal effect of foreign aid on economic growth remains on the other hand positive and highly statistically significant.[10]

5 Further Issues

5.1 Cross-Country Parameter Heterogeneity

The log-log difference specification (see equation (2)) implies that, cross-country differences in the effect that a change in the level of foreign aid has on the level of GDP per capita are differenced out. However, it is likely that also the elasticity effect of foreign aid on GDP per capita growth differs across countries (as would for example be suggested by a standard Solow-Swan growth model; see Section 5.2 below). To check whether parameter heterogeneity leads to a bias in the estimated average effect, I use the mean-group estimator developed by Pesaran and Smith (1995) that computes estimates country-by-country and then takes a linear average of the obtained coefficients. Figure 1 plots the kernel density function of the country-specific slope estimates that are obtained from using as an instrumental variable the residual variation in foreign aid that is not driven by economic growth.[11] The mean value of the country-specific slope estimates is 0.11, and thus matches closely the estimate of the average marginal effect reported in column (1) of Table 3 from the homogenous panel fixed effects model.

Beyond providing an important robustness check on the average marginal effect obtained from the homogenous panel fixed effects model, the country-specific slope estimates provide useful information on the extent to which the

[9] See, for example, Trumball and Wall (1994), or Alesina and Dollar (2000).

[10] Panel B of Table 4 shows that increases in countries' Polity2 scores are associated with significant increases in foreign aid. While there is no significant contemporaneous effect of economic growth on the Polity2 score when using rainfall and the international commodity price index as instruments for GDP growth (results not shown), the correlation in Panel B is unlikely to reflect the true causal effect that foreign aid has on economic growth. Barro and Lee (2005) and Djankov et al. (2008), for example, provide evidence that foreign aid can have adverse effects on countries' political institutions. In this case, the estimate in Panel B of Table 4 reflects a lower bound on the true causal effect that political change has on foreign aid.

[11] The reported estimates in Figure 1 are based on the largest possible sample (47 countries during 1960-2000).

effect of foreign aid on economic growth varies across countries. The interquartile range of the country-specific slope estimates is [0.05,0.15], with a sample minimum (maximum) of -0.02 (0.35). Hence, there is quite a bit of variation in the marginal effect that foreign aid has on economic growth across countries, which raises the interesting policy question of what determines this cross-country variation.

One explanation for the cross-country variation in the marginal effect of foreign aid on economic growth are cross-country differences in economic policies. Burnside and Dollar (2000) argued that the marginal effect of foreign aid on economic growth is particularly high in countries where policy-induced distortions to economic activity are relatively small because in these countries aid is more likely to be invested.[12] Figure 2 examines this claim empirically by plotting the country-specific slope estimates against the Burnside and Dollar (2000) policy index that captures cross-country differences in trade policy, inflation, and budget balance. The scatter plot shows a positive relationship between the country-specific slope estimates and the (period-average) BD policy index. Using the bootstrap to take into account the relatively small number of observations, a bivariate regression yields a coefficient on the BD policy index of 0.04 that has a t-value of 1.86. Hence, this is supportive evidence for the Burnside and Dollar claim that foreign aid is particularly effective in stimulating economic growth in countries with good macroeconomic policies.

Dalgaard et al. (2004) found that aid is significantly less effective in the tropics. As noted by Rajan and Subramanian (2008), there is little theoretical reason for why one would expect a systematically smaller effect of foreign aid on economic growth in countries which are located in the tropics. Figure 3 shows that regressing the country-specific slope estimates on the share of tropical terrain yields a negative, but statistically insignificant relationship.

Another argument for cross-country heterogeneity in the marginal effect of foreign aid on economic growth, that has been popular in both academic and policy circles, are financing constraints (see e.g. Sachs, 2005). Domestic and, in particular, rural financial markets are often ill-functioning (or simply nonexistant) in many of the LDCs so that high return projects go unrealized because (rural) investors fail to obtain finance for their projects. An aid inflow may have a high return if it successfully targets high return projects and eases financing constraints in the (rural) financial markets. Figure 4, Panels A-C explore the role of such financing constraints by plotting the country-specific slope estimates against various indicators proxying the severity of financial market imperfections. Panel A plots the relationship between the country-specific slope estimates and the World Bank credit information index that captures the availability of credit information from either a public registry or a private bureau to facilitate lending decisions. Panel B plots the relationship between the country-specific slope estimates and the percentage share of individuals and firms listed in a public or private credit registry with current information on repayment history, unpaid debts, or credit outstanding. And, to capture that credit market

[12]For a critique, see Easterly et al. (2004).

imperfections are often most severe in rural areas of developing countries Panel C plots the relationship between the country-specific slope estimates and the percentage share of the population living in rural regions. The main conclusion is that the marginal effect of foreign aid on economic growth is significantly increasing in these proxies for cross-country differences in the severity of financing constraints. Hence, the common argument in favor of foreign aid – the financing problem – finds support in the data.

A strand of the aid effectiveness literature has argued that there exists a political economy resource curse of foreign aid on economic growth: in countries with multiple powerful groups aid inflows may lead to costly rent-seeking activity (e.g. Svensson, 2000; Reinikka and Svensson, 2004). Figure 5, Panels A and B explore this channel by plotting the relationship between the country-specific slope estimates and two measures that capture countries' ethnic fragmentation. Panel A plots the relationship using an index of ethnic fractionalization and Panel B plots the relationship using an index of ethnic polarization.[13] Both figures show a downward sloping relationship. Statistically the relationship is however only significant at conventional confidence levels for the measure of ethnic fractionalization, which may suggest that aid inflows can be a curse primarily due to the common pool problem, rather than because they directly increase the likelihood of civil conflict.[14]

5.2 Comparison of IV Estimates to the Predicted Effect from a Solow-Swan Growth Model

A useful way to check whether the instrumental variables estimates of the average marginal effect are plausible not only in sign but also in size is to draw on the first-order approximation of the effect that a change in the investment rate has on the output growth rate in a simple but standard Solow-Swan growth model. The first-order approximation yields that a 1 percent increase in the investment rate increases the output growth rate by $\beta\frac{\alpha}{1-\alpha}$ percentage points, where β is the convergence rate and α the capital-output elasticity.[15] If part of the foreign aid is used to finance domestic investment, the predicted growth rate effect (in percentage points) of a 1 percent increase in the share of aid in GDP is $\beta\frac{\alpha}{1-\alpha}$ times the marginal elasticity effect that foreign aid has on investment.[16]

[13] The fractionalization index increases with the number of groups, while the polarization index is maximized when there are two groups which are of equal size. Both indices are between 0 and 1, with larger values denoting more fractionalization (polarization). For a discussion of conceptual differences between polarization and fractionalization indices with an application to the conflict literature, see Montalvo and Reynal-Querol (2005).

[14] In fact, recent research by De Ree and Nillesen (2009) shows that an increase in foreign aid is associated with a significant decrease in the likelihood of civil conflict.

[15] See for example Barro and Sala-i-Martin (2003).

[16] Arellano et al. (2008) show in a DSGE model, where consumers are modelled to perfectly smooth consumption over time, that whether an aid inflow increases investment depends on the persistence of the aid shock. A fully permanent aid shock increases consumption, with little effect on investment – a result that follows from the Permanent Income Hypothesis. It is questionable however whether the Permanent Income Hypothesis is a good approximation

Table 5 shows that on average this marginal elasticity effect of foreign aid on investment is around 0.15-0.3 percent.

To make the instrumental variables estimates comparable therefore to the predictions of the Solow-Swan growth model, and in order to obtain an estimate of the convergence rate β, Table 6 shows that there continues to be a positive and statistically significant effect of foreign aid on the growth rate of GDP per capita when controlling for convergence effects in the level of GDP per capita and using instead of the log of foreign aid the log of the share of foreign aid in GDP. The system-GMM estimation (Blundell and Bond, 1998) for these dynamic regressions, where the foreign aid share is instrumented by the variation in the aid to GDP ratio that is not driven by GDP per capita growth, yields estimates on the log of the aid to GDP ratio that range between 0.05 to 0.07 for the 1960-2000 period and between 0.03 to 0.06 for the 1970-2000 period. Statistically, these estimates are significant at the 5 percent level at least. The average annual convergence rate in the sample is estimated to be around 5 to 10 percent per annum.

With these estimates in hand, it is now possible to compare the instrumental variables estimates of the average effect that foreign aid has on economic growth to the quantitative prediction from the Solow-Swan growth model. A typical value used in the empirical growth literature for α, that is reasonable for a Solow-Swan growth model with investment in both, physical and human capital, is 2/3 (see e.g. Durlauf et al., 2005).[17] Hence, the predicted average effect of a 1 percent change in the aid to GDP ratio on the output growth rate is around 0.02 to 0.06 percentage points. This is in line with the instrumental variables estimates reported in Table 6, which range between 0.03 to 0.07.

5.3 Level Effects vs. Growth Effects

The Solow-Swan growth model predicts that a permanent increase in foreign aid affects GDP per capita growth along the transition to the new steady-state. However, due to the assumption that there are decreasing returns to scale in capital the Solow-Swan growth model predicts that an increase in foreign aid has a level effect but not a long-run growth effect. The empirical results so far are consistent with both a level effect and a growth effect. This is because the

of consumption choices in the LDCs. Empirically there exists evidence that for the LDCs the Permanent Income Hypothesis is not a good approximation, mainly because of financing constraints (see, for example, Deaton, 1992).

[17]Note that in the Solow-Swan growth model there exists a tight relationship between the convergence rate β and the output-capital elasticity α. In particular, it holds that $\beta = (1-\alpha)(n+g+d)$, where n and g are the population and TFP growth rates respectively, and d is the depreciation rate (e.g. Barro and Sala-i-Martin, 2003). The average sample population growth rate is about 2.5 percent and a reasonable value for annual TFP growth is about 1 to 2 percent. An α of 2/3 and an estimated convergence rate of 5 to 10 percent would therefore require a depreciation rate of the capital stock of between 10 to 25 percent per annum. For the LDCs, where weather conditions are often extreme, this may not be unreasonable. Bu (2006), for example, presents firm data evidence for Ghana, Ivory Coast, Kenya, and Zimbabwe with average depreciation rates on fixed assets (resp. machinery and equipment) that range between 10 to 20 percent (resp. 15 to 40 percent).

first-difference specification that relates the log-change in GDP per capita to the log-change in foreign aid has an analogous level form representation where the log of GDP per capita is related to the log of foreign aid.

To examine whether beyond a level effect an increase in foreign aid has also an effect on the long-run GDP per capita growth rate, I include as an additional right-hand-side regressor in the growth equation the log of the level of foreign aid. This approach follows the empirical growth literature that has tested for long-run growth effects of investment.[18] In the growth equation, the estimated coefficient on the level of foreign aid reflects the effect that foreign aid has on the long-run GDP per capita growth rate while the estimated coefficient on the log-change of foreign aid reflects the effect that foreign aid has on the level of GDP per capita.

Table 7 reports the estimates for the largest possible sample of 47 countries during the 1960-2000 period. The main result is that the estimated coefficient on the level of foreign aid is statistically insignificant and quantitatively small. On the other hand, the estimated coefficient on the log-change of foreign aid is positive, highly statistically significant, and quantitatively large. Table 7 therefore shows that a permanent increase in foreign aid has a significant positive effect on the level of GDP per capita but an insignificant effect on the long-run GDP per capita growth rate. This result holds for both the static and the dynamic panel data model (columns (1) and (2)). And, it also holds in a distributed lag model where additional lags of foreign aid are included on the right-hand side of the estimating equation (columns (3)-(5)). In particular, the distributed lag estimates (columns (3)-(5)) show that foreign aid has a positive and statistically significant effect on GDP per capita growth on impact, and that the lagged effects are quantitatively smaller in size. The sum of the coefficients on the contemporaneous and lagged log-changes of foreign aid is positive and significantly different from zero at the 1 percent level. Hence, a permanent increase in foreign aid has a significant positive long-run effect on the level of GDP per capita. This result is consistent with the neoclassical Solow-Swan growth model where part of the foreign aid is used to finance domestic investment.

6 Conclusion

This paper showed as a first main result that increases in per capita GDP growth of aid recipient countries are associated with a significant decrease in foreign aid. Specifically, the instrumental variables estimates yielded that a 1 percentage point increase in per capita GDP growth reduced foreign aid by more than 4 percent on average. This finding is consistent with the stylized cross-country fact that as countries grow richer they rely less on foreign aid. It is also consistent with donor countries acting as Good Samaritans.

The paper's finding of a quantitatively large, negative effect of economic growth on foreign aid bears an important implication for empirical research on aid effectiveness: OLS estimates which serve as a natural benchmark of

[18] See for example Bond et al. (2010) and the references cited therein.

comparison to are biased against finding a significant positive average effect of foreign aid on economic growth. Hence, insignificant estimates of the effect of foreign aid on economic growth should be viewed with skepticism – they may just be a consequence of an inadequately addressed negative and quantitatively large simultaneity bias.

As a second main finding, the paper showed that after the large, negative response of foreign aid to per capita GDP growth is accounted for that foreign aid did indeed have a statistically significant positive effect on per capita GDP growth. This finding contrasts to recent empirical papers that have failed to find a significant positive average effect of foreign aid on economic growth. Because there is a strong tendency at the macroeconomic level for foreign aid to decrease as per capita GDP growth of aid recipient countries increases, the cards are stacked in empirical research against finding a significant positive average effect of foreign aid on economic growth. Hence, if the reverse causality running from higher per capita GDP to less foreign aid is not properly addressed, the researcher may fail to find a significant positive average effect of foreign aid on economic growth and possibly conclude that foreign aid does not have a systematic positive average effect on per capita GDP growth.

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8 Technical Appendix

8.1 Adjustment of the Simultaneity Bias

This appendix shows formally why an instrumental variables estimator, that uses the variation in foreign aid which is not driven by economic growth as an instrumental variable for foreign aid is immune to the simultaneity bias. For clarity, let us consider the simplest possible simultaneous equation model:

$$GDP = \gamma_1 Aid + u \tag{5}$$

$$Aid = \gamma_2 GDP + e \tag{6}$$

The probability limit of the OLS estimator of γ_1 in equation (5) is:

$$\gamma_1^{OLS} = \gamma_1 + \frac{cov(Aid, u)}{var(Aid)} \tag{7}$$

Substituting equation (5) into (6), and rearranging yields the equilibrium condition:

$$Aid = \frac{1}{1 - \gamma_1\gamma_2}(\gamma_2 u + e) \tag{8}$$

Hence, by substitution of equation (8) into (7) yields

$$\gamma_1^{OLS} = \gamma_1 + \frac{\gamma_2}{1 - \gamma_1\gamma_2}\frac{var(u)}{var(Aid)} + \frac{1}{1 - \gamma_1\gamma_2}\frac{cov(e, u)}{var(Aid)}, \tag{9}$$

where the second term on the right-hand side of equation (9) captures the simultaneity bias that arises if $\gamma_2 \neq 0$ in equation (6), and the third term captures the omitted variables bias.[19]

Suppose now that one is able to obtain a consistent estimate of γ_2 in equation (6).[20] Using this consistent estimate, one can construct an aid series Aid^* that is adjusted for the endogenous response (i.e. $Aid^* = Aid - \gamma_2 GDP$) and use this variable as an instrument for the original aid variable Aid in equation (5). The probability limit of this IV estimator is

$$\gamma_1^{IV} = \frac{cov(Aid^*, GDP)}{cov(Aid^*, Aid)} = \gamma_1 + \frac{cov(Aid^*, u)}{cov(Aid^*, Aid)} = \gamma_1 + \frac{cov(e, u)}{cov(e, Aid)}. \tag{10}$$

Hence, the IV estimator that uses the endogeneity adjusted aid series Aid^* as an instrument for Aid does not suffer from the simultaneity bias.

[19] To see that the third term in equation (9) captures the omitted variables bias, set $\gamma_2 = 0$. In this case $Aid = u$ and the probability limit of the OLS estimator is simply $\gamma_1 + \frac{cov(Aid, e)}{var(Aid)}$.

[20] Of course, this can only be done by having a valid instrument for GDP in equation (6). OLS cannot provide in equation (6) a consistent estimate for γ_2 if in equation (5) $\gamma_1 \neq 0$.

8.2 Size of the Omitted Variables Bias

Regarding the size of the omitted variables bias of the IV estimator (OVB^{IV}) that arises if $cov(e,u) \neq 0$, note that the second term in equation (10) simplifies to

$$OVB^{IV} = (1 - \gamma_1\gamma_2)\frac{cov(e,u)}{var(e) + \gamma_2 cov(e,u)}. \tag{11}$$

The third term in equation (9) that captures the omitted variables bias of the least squares estimator (OVB^{IV}) simplifies to

$$OVB^{LS} = (1 - \gamma_1\gamma_2)\frac{cov(e,u)}{var(e) + \gamma_1^2 var(u) + 2\gamma_2 cov(e,u)}. \tag{12}$$

Depending on $\gamma_1^2 var(u)$ and the sign and size of $\gamma_2 cov(e,u)$, the omitted variables bias of the IV estimator may, therefore, be smaller or larger than the omitted variables bias of the least squares estimator.

9 Data Appendix

List of Countries and Summary Statistics

Country	Aid/GDP	Agri/GDP	Export/GDP	Country	Aid/GDP	Agri/GDP	Export/GDP
Afghanistan	0.4	58	18	Madagascar	1.8	62	20
Bangladesh	0.8	73	8	Malawi	3.4	37	22
Benin	2.1	19	29	Maldives	3.1	28	89
Bhutan	3.0	9	26	Mali	3.3	27	20
Burkina Faso	2.7	32	10	Mauritania	7.3	38	41
Burundi	2.3	78	9	Mozambique	2.8	62	10
Cambodia	1.1	20	20	Nepal	0.9	28	13
Cameroon	0.9	19	22	Niger	2.5	26	19
Cape Verde	6.9	17	19	Rwanda	3.0	65	8
Chad	2.3	38	23	Samoa	5.5	35	31
Comoros	4.3	62	14	Sao Tome &Principe	15.7	42	24
Djibouti	7.2	59	43	Senegal	3.0	42	26
Congo, Rep. of	2.2	31	58	Sierra Leone	2.0	43	19
Eq. Guinea	4.1	12	55	Solomon Islands	4.9	23	47
Eritrea	5.8	73	15	Somalia	4.6	70	4
Ethiopia	1.6	31	11	Sudan	1.9	50	10
Gambia	4.9	59	7	Tanzania	3.9	36	17
Guinea	0.9	49	26	Togo	2.6	57	33
Guinea-Bissau	11.0	51	16	Uganda	1.8	55	11
Haiti	1.4	58	19	Vanuatu	7.7	11	46
Kiribati	13.7	47	28	Yemen	2.2	34	27
Laos	2.4	7	15	Zambia	4.1	31	34
Lesotho	3.3	78	24	Zimbabwe	0.5	32	28
Liberia	7.6	27	45	Average	3.9	42	24

Note: The table lists for each country the sample average share of foreign aid in GDP, the sample average share of agricultural value added in GDP, and the sample average share of exports in GDP (Source: WDI 2009). All values have been multiplied by 100.

Table 1: The Effect of Economic Growth on Foreign Aid
(Baseline IV Estimates)

	$\Delta\ln(y)$	$\Delta\ln(aid)$						
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	LS	LS	LS	2SLS	2SLS	2SLS	2SLS	2SLS
$\Delta\ln(y)$			-0.37**	-4.47**	-4.09*	-3.92*	-6.04***	-4.73**
			(-2.48)	[0.04]	[0.07]	[0.07]	[0.01]	[0.05]
$\Delta\ln(ComPI)$	0.35***	-2.12**			-0.69	-0.74		
	(3.45)	(-2.40)			(-0.48)	(-0.54)		
$\Delta\ln(Rain)$	0.18***	-0.69*			0.03		0.37	
	(3.09)	(-1.76)			(0.31)		(0.46)	
$\Delta\ln(Rain)^2$	-0.01***	0.05*				0.00	-0.02	
	(-2.93)	(1.75)				(0.32)	(-0.42)	
Hansen J, p-value	.	.	.	0.82	.	.	.	0.48
First Stage F-stat	.	.	.	9.32	8.56	9.54	11.94	15.85
Country FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations	1550	1550	1550	1550	1550	1550	1550	1316
Countries	47	47	47	47	47	47	47	47

Note: The method of estimation in columns (1)-(3) is least squares; columns (4)-(8) two-stage least squares. t-values (in parentheses) are reported below the least-squares estimates; below two-stage least squares estimates p-values [in square brackets] are reported based on the Anderson-Rubin test of statistical significance. All test statistics are based on Huber robust standard errors that are clustered at the country level. The instrumental variables in columns (4)-(8) are the log-changes in the commodity price index, rainfall, and rainfall squared. Column (8) shows two-stage least squares estimates excluding all those country-year observations where countries experienced a drought year (as reported by the Catholique de Louvain's Emergency Disaster database). *Significantly different from zero at 90 percent confidence, ** 95 percent confidence, *** 99 percent confidence.

Table 2. The Effect of Economic Growth on Foreign Aid
(Robustness of IV Estimates to Different Time Periods, Outliers, and Balanced Panel)

	Δln(aid)				
	(1)	(2)	(3)	(4)	(5)
	1960-2000 Balanced Panel	1960-2000 Balanced Panel & Excluding Outliers	1970-2000 All LDCs	1970-2000 Balanced Panel	1970-2000 Balanced Panel & Excluding Outliers
Δln(y)	-5.47***	-5.92***	-4.19*	-5.03**	-6.07***
	[0.00]	[0.00]	[0.09]	[0.04]	[0.00]
Hansen J, p-value	0.49	0.54	0.91	0.96	0.68
First Stage F-stat	7.76	10.08	8.14	10.08	8.60
Country FE	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes
Observations	819	811	1341	1170	1150
Countries	21	21	47	39	39

Note: The method of estimation is two-stage least squares. The p-values [in square brackets] are based on the Anderson-Rubin test of statistical significance. All test statistics are based on Huber robust standard errors that are clustered at the country level. The instrumental variables are the log-changes in the commodity price index, rainfall, and rainfall squared. *Significantly different from zero at 90 percent confidence, ** 95 percent confidence, *** 99 percent confidence.

Table 3: The Effect of Foreign Aid on Economic Growth
(Baseline IV Estimates)

$\Delta\ln(y)$

Panel A: IV-2SLS Estimates

	(1)	(2)	(3)	(4)	(5)	(6)
	1960-2000 All LDCs	1960-2000 Balanced Panel	1960-2000 Balanced Panel & Excluding Outliers	1970-2000 All LDCs	1970-2000 Balanced Panel	1970-2000 Balanced Panel & Excluding Outliers
$\Delta\ln(\text{aid})$	0.12***	0.23***	0.23***	0.15***	0.18***	0.23***
	(3.36)	(4.39)	(6.26)	(3.46)	(2.72)	(7.18)
First-Stage F-stat	92.03	55.77	106.85	78.16	31.52	141.78
Country FE	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes
Observations	1550	819	811	1341	1170	1150
Countries	47	21	21	47	39	39

Panel B: LS Estimates

	(1)	(2)	(3)	(4)	(5)	(6)
	1960-2000 All LDCs	1960-2000 Balanced Panel	1960-2000 Balanced Panel & Excluding Outliers	1970-2000 All LDCs	1970-2000 Balanced Panel	1970-2000 Balanced Panel & Excluding Outliers
$\Delta\ln(\text{aid})$	-0.01**	-0.01	-0.00	-0.01*	-0.01**	0.00
	(-2.25)	(-0.82)	(-0.25)	(-1.95)	(-2.47)	(0.02)
Country FE	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes
Observations	1550	819	811	1341	1170	1150
Countries	47	21	21	47	39	39

Note: The method of estimation in Panel A is two-stage least squares; Panel B least squares. t-values (in brackets) are based on Huber robust standard errors that are clustered at the country level. The instrumental variable for the two-stage least squares estimation in Panel A is the foreign aid series that is adjusted for the reverse effect that per capita GDP growth has on foreign aid. See Section 2.2 for a detailed explanation of this estimation strategy. All regressions control for the log-changes in the commodity price index, rainfall, and rainfall squared (estimates not shown). *Significantly different from zero at 90 percent confidence, ** 95 percent confidence, *** 99 percent confidence.

Table 4: The Effect of Foreign Aid on Economic Growth
(Controlling for Changes in Political Institutions)

Panel A: Dependent Variable is $\Delta\ln(y)$

	(1)	(2)	(3)	(4)	(5)	(6)
	1960-2000 All LDCs	1960-2000 Balanced Panel	1960-2000 Balanced Panel & Excluding Outliers	1970-2000 All LDCs	1970-2000 Balanced Panel	1970-2000 Balanced Panel & Excluding Outliers
$\Delta\ln(\text{aid})$	0.18***	0.25***	0.24***	0.16***	0.22***	0.22***
	(4.64)	(4.14)	(5.57)	(4.31)	(3.37)	(7.37)
$\Delta(\text{polity2})$	0.00	0.00	-0.01	0.00	-0.00	0.00
	(0.52)	(0.02)	(-0.18)	(0.40)	(-0.95)	(0.09)
First-Stage F-stat	122.68	43.17	76.43	181.13	20.79	174.52
Country FE	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes
Observations	1265	800	793	1093	984	972
Countries	39	20	20	39	34	34

Panel B: Dependent Variable is $\Delta\ln(\text{aid})$

	(1)	(2)	(3)	(4)	(5)	(6)
	1960-2000 All LDCs	1960-2000 Balanced Panel	1960-2000 Balanced Panel & Excluding Outliers	1970-2000 All LDCs	1970-2000 Balanced Panel	1970-2000 Balanced Panel & Excluding Outliers
$\Delta\ln(y)$	-3.82**	-5.84***	-6.29***	-2.88*	-4.55***	-5.34**
	(-2.34)	(-3.41)	(3.97)	(-1.80)	(-2.81)	(-2.46)
$\Delta(\text{polity2})$	0.02**	0.01	0.02	0.02**	0.02**	0.01
	(2.20)	(1.09)	(1.34)	(2.46)	(2.01)	(1.02)
First-Stage F-stat	6.78	7.18	8.58	5.90	6.91	4.94
Country FE	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes
Observations	1265	800	793	1093	984	972
Countries	39	20	20	39	34	34

Note: The method of estimation is two-stage least squares. The instrumental variable for the two-stage least squares estimation in Panel A is the foreign aid series that is adjusted for the reverse effect that per capita GDP growth has on foreign aid. See Section 2.2 for a detailed explanation of this estimation strategy. The instrumental variables for the two-stage least squares estimation in Panel B are the log-changes in the commodity price index, rainfall, and rainfall squared. *Significantly different from zero at 90 percent confidence, ** 95 percent confidence, *** 99 percent confidence.

Table 5: The Effect of Foreign Aid on Economic Growth
(Investment Response)

	Δln(i)					
	(1)	(2)	(3)	(4)	(5)	(6)
	1960-2000 All LDCs	1960-2000 Balanced Panel	1960-2000 Balanced Panel & Excluding Outliers	1970-2000 All LDCs	1970-2000 Balanced Panel	1970-2000 Balanced Panel & Excluding Outliers
Δln(aid)	0.14***	0.30***	0.31***	0.15**	0.17**	0.24***
	(2.70)	(3.00)	(2.70)	(2.30)	(2.22)	(3.30)
First-Stage F-Statistic	92.03	55.77	106.85	78.16	31.52	141.78
Country FE	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes
Observations	1550	819	811	1341	1170	1150
Countries	47	21	21	47	39	39

Note: The method of estimation is two-stage least squares. t-values (in brackets) are based on Huber robust standard errors that are clustered at the country level. The instrumental variable for the two-stage least squares estimation is the foreign aid series that is adjusted for the reverse effect that per capita GDP growth has on foreign aid. See Section 2.2 in the paper for a detailed explanation of this estimation strategy. All regressions control for the log-changes in the commodity price index, rainfall, and rainfall squared (estimates not shown). *Significantly different from zero at 90 percent confidence, ** 95 percent confidence, *** 99 percent confidence.

Table 6: The Effect of Foreign Aid on Economic Growth
(Using the Share of Aid in GDP and Controlling for Convergence Effects)

	$\Delta\ln(y)$					
	(1)	(2)	(3)	(4)	(5)	(6)
	1960-2000 All LDCs	1960-2000 Balanced Panel	1960-2000 Balanced Panel & Excluding Outliers	1970-2000 All LDCs	1970-2000 Balanced Panel	1970-2000 Balanced Panel & Excluding Outliers
$\Delta\ln(aid/y)$	0.05***	0.07***	0.07***	0.06***	0.06**	0.03**
	(3.25)	(3.76)	(3.10)	(3.11)	(2.28)	(1.99)
L.ln(y)	-0.06**	-0.12***	-0.10***	-0.06*	-0.05**	-0.05
	(-1.98)	(-5.05)	(-4.89)	(-1.87)	(-2.24)	(-1.20)
First-Stage F-Statistic	49.19	25.75	47.84	37.64	15.34	63.99
Country FE	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes
Observations	1550	819	811	1341	1170	1150
Countries	47	21	21	47	39	39

Note: The method of estimation is system-GMM (Blundell and Bond, 1998). t-values (in brackets) are based on Huber robust standard errors that are clustered at the country level. The instrumental variable for the share of foreign aid in GDP is the share of foreign aid in GDP that is adjusted for the reverse effect that per capita GDP growth has on the share of foreign aid in GDP. See Section 2.2 in the paper for a detailed explanation of this estimation strategy. All regressions control for the log-changes in the commodity price index, rainfall, and rainfall squared (estimates not shown). *Significantly different from zero at 90 percent confidence, ** 95 percent confidence, *** 99 percent confidence.

Table 7: The Effect of Foreign Aid on Economic Growth
(Level Effects vs. Growth Effects)

			Δln(y)		
	(1)	(2)	(3)	(4)	(5)
	1960-2000 All LDCs	1960-2000 All LDCs	1960-2000 All LDCs	1960-2000 All LDCs	1960-2000 All LDCs
ln(aid)	0.01 (1.32)	-0.01 (-0.87)	-0.01 (-1.45)	-0.01 (-1.03)	-0.01 (-1.04)
Δln(aid)	0.11*** (3.49)	0.12*** (3.47)	0.14*** (3.03)	0.17*** (2.92)	0.18*** (3.03)
L.Δln(y)		0.83*** (7.71)	0.82*** (7.07)	0.82*** (5.44)	0.79*** (4.99)
L.Δln(aid)			0.03** (2.04)	0.03 (1.41)	0.04 (1.52)
L2.Δln(aid)				0.01 (0.18)	0.01 (1.10)
L3.Δln(aid)					0.01 (1.31)
Sum of Coefficients on Δln(aid)			0.17*** (2.95)	0.20*** (2.64)	0.24*** (3.00)
First-Stage F-Statistic	7.54	7.27	6.84	6.79	6.34
Country FE	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes
Observations	1550	1497	1488	1440	1425
Countries	47	47	47	47	47

Note: The method of estimation in column (1) is two-stage least squares; columns (2)-(5) system-GMM (Blundell and Bond, 1998). t-values (in brackets) are based on Huber robust standard errors that are clustered at the country level. The instrumental variable for foreign aid is the foreign aid series that is adjusted for the reverse effect that per capita GDP growth has on foreign aid. The instrumental variable for lagged GDP per capita growth is the second and third lag of GDP per capita. All regressions control for the log-changes in the commodity price index, rainfall, and rainfall squared (estimates not shown). *Significantly different from zero at 90 percent confidence, ** 95 percent confidence, *** 99 percent confidence.

Appendix Table A1: Weak IV Estimators

	Δln(aid)					
	(1)	(2)	(3)	(4)	(5)	(6)
	1960-2000 All LDCs	1960-2000 Balanced Panel	1960-2000 Balanced Panel & Excluding Outliers	1970-2000 All LDCs	1970-2000 Balanced Panel	1970-2000 Balanced Panel & Excluding Outliers
Δln(y) (Fuller 1)	-4.20** (-2.20)	-5.78*** (-3.46)	-6.19*** (-4.25)	-3.87** (-2.00)	-4.58** (-2.61)	-6.24*** (-2.65)
Δln(y) (Fuller 4)	-3.31** (-2.31)	-4.57*** (-4.12)	-4.82*** (-4.31)	-2.96** (-2.15)	-3.60*** (-2.86)	-4.63*** (-3.14)
Δln(y) (GMM CUE)	-4.85** (-2.35)	-5.98*** (-5.02)	-5.84*** (-5.48)	-4.25** (-2.03)	-4.95** (-2.52)	-7.66*** (-3.19)
First-Stage F-Statistic	9.32	7.76	10.03	8.14	10.08	8.18
Country FE	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes
Observations	1550	819	811	1341	1170	1150
Countries	47	21	21	47	39	39

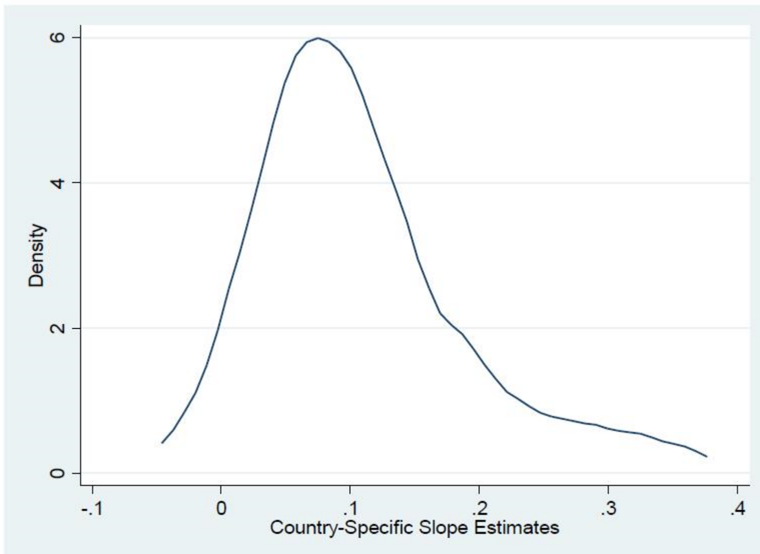
Note: The method of estimation in the first (second) row is the Fuller (1977, *Econometrica*) modified LIML estimator, with an alpha constant set equal to 1 (4). In the third row the method of estimation is the Hansen et al. (1996, *Journal of Business and Economic Statistics*) continuously updated GMM estimator. t-values (in parentheses) are based on Huber robust standard errors that are clustered at the country level. The instrumental variables are the log-changes in the commodity price index, rainfall, and rainfall squared. *Significantly different from zero at 90 percent confidence, ** 95 percent confidence, *** 99 percent confidence.

Appendix Table A2: Results Using Other Datasets

	Δln(y)		
	Panel A: IV-2SLS Estimates		
	(1)	(2)	(3)
Dataset From:	Burnside and Dollar (2000)	Easterly et al. (2004)	Roodman (2007)
Δln(aid)	0.19*** (3.88)	0.10*** (4.31)	0.16*** (3.04)
First-Stage F-stat	77.71	227.72	58.96
Country FE	Yes	Yes	Yes
Year FE	Yes	Yes	Yes
Observations	431	537	728
Countries	94	111	124
	Panel B: LS Estimates		
	(1)	(2)	(4)
Dataset From:	Burnside and Dollar (2000)	Easterly et al. (2004)	Roodman (2007)
Δln(aid)	-0.04** (-2.27)	-0.03** (-2.44)	-0.05*** (-4.19)
Country FE	Yes	Yes	Yes
Year FE	Yes	Yes	Yes
Observations	431	537	728
Countries	94	111	124

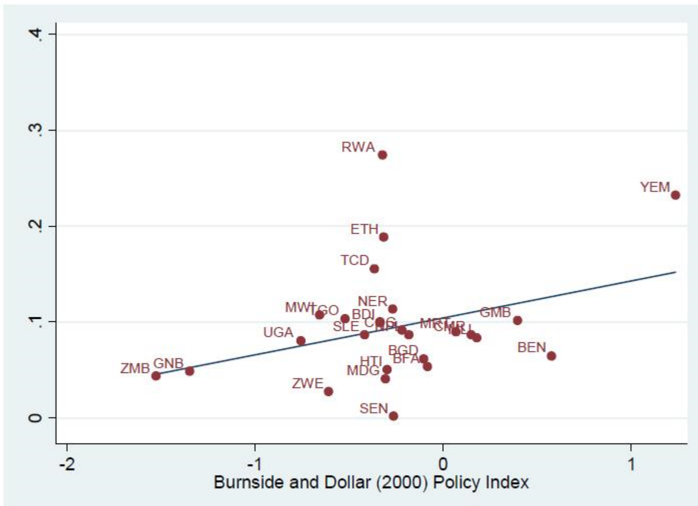
Note: The method of estimation in Panel A is two-stage least squares; Panel B least squares. t-values (in brackets) are based on Huber robust standard errors that are clustered at the country level. The instrumental variable for the two-stage least squares estimation in Panel A is the foreign aid series that is adjusted for the reverse causal effect that per capita GDP growth has on foreign aid, using a structural coefficient of -4 that is a lower bound of the estimates obtained in Tables 1 and 2. *Significantly different from zero at 90 percent confidence, ** 95 percent confidence, *** 99 percent confidence.

Figure 1. Distribution of Country-Specific Slope Estimates



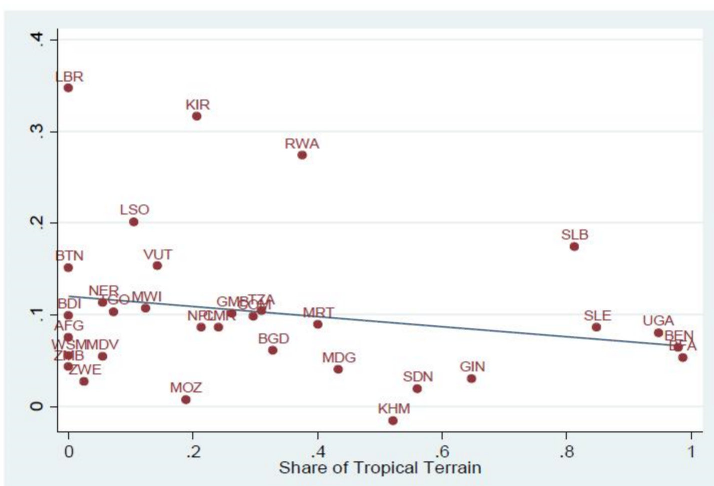
Note: The figure shows the density function of the country-specific slope estimates that are obtained by applying the Pesaran and Smith (1995) mean-group estimator and instrumenting the foreign aid series by the residual variation in foreign aid that is not driven by economic growth. The density function is estimated using an Epanechnikov kernel, with bandwidth selection based on Silverman's rule of thumb.

Figure 2. The Role of Macroeconomic Policies



Note. The figure shows the relationship between the country-specific slope estimates (reported in Figure 1) and the Burnside and Dollar (2000) policy index. The bootstrapped slope coefficient (s.e.) of the fitted regression line is 0.04 (0.02).

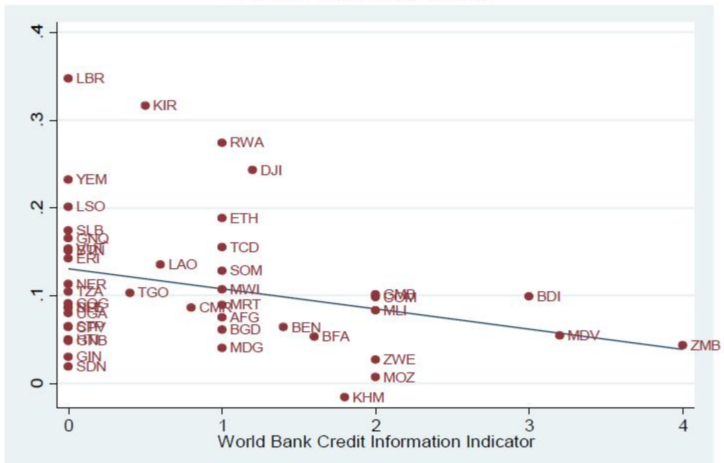
Figure 3. The Role of Tropical Terrain



Note: The figure shows the relationship between the country-specific slope estimates (reported in Figure 1) and the share of tropical terrain. The bootstrapped slope coefficient (s.e.) of the fitted regression line is -0.05 (0.04).

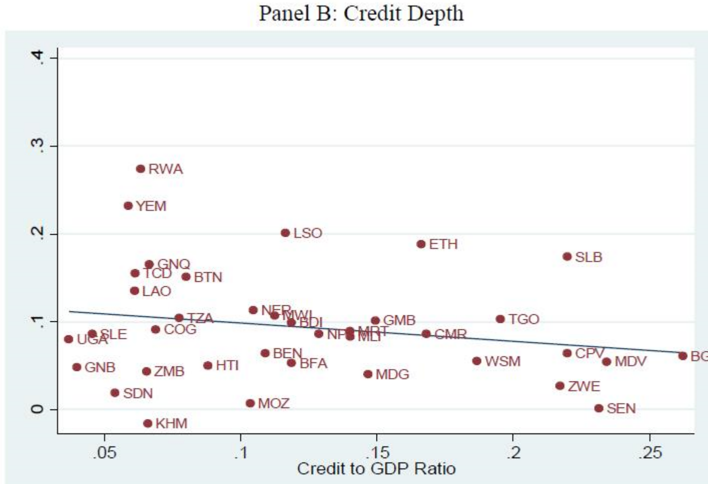
Figure 4. The Role of Financing Constraints

Panel A: Credit Information



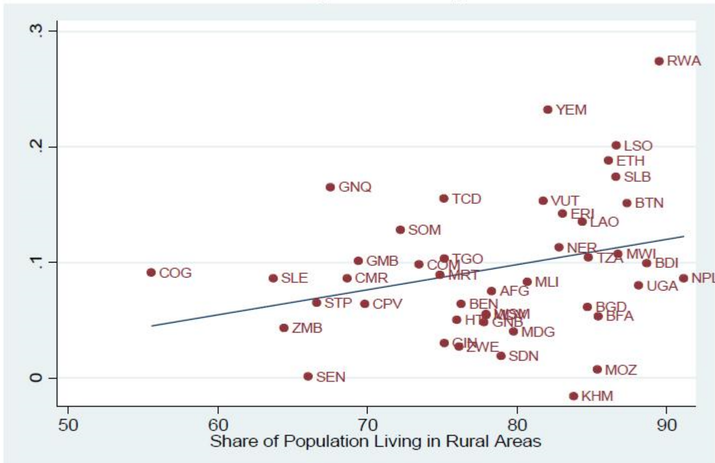
Note: The figure shows the relationship between the country-specific slope estimates (reported in Figure 1) and World Bank credit information indicator. Higher values of the credit information indicator represent better credit information. The bootstrapped slope coefficient (s.e.) of the fitted regression line is -0.02 (0.01).

Panel B: Credit Depth



Note: The figure shows the relationship between the country-specific slope estimates (reported in Figure 1) and the ratio of public and private credit to GDP. The bootstrapped slope coefficient (s.e.) of the fitted regression line is -0.21 (0.16).

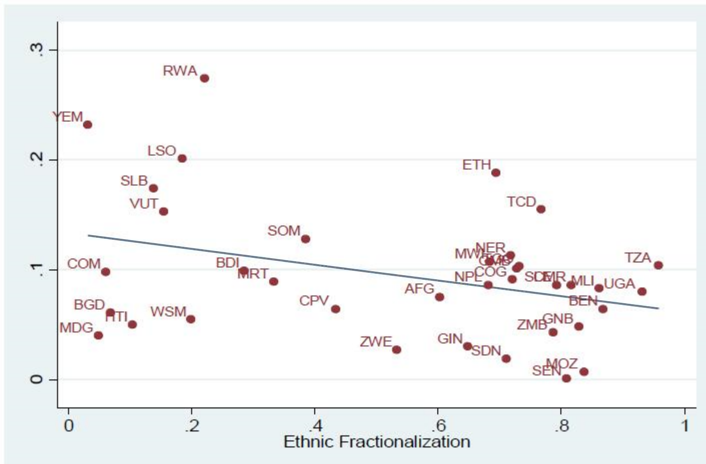
Panel C: Percent of Population Living in Rural Areas



Note: The figure shows the relationship between the country-specific slope estimates (reported in Figure 1) and the percentage share of the population living in rural regions. The bootstrapped slope coefficient (s.e.) of the fitted regression line is 0.002 (0.001).

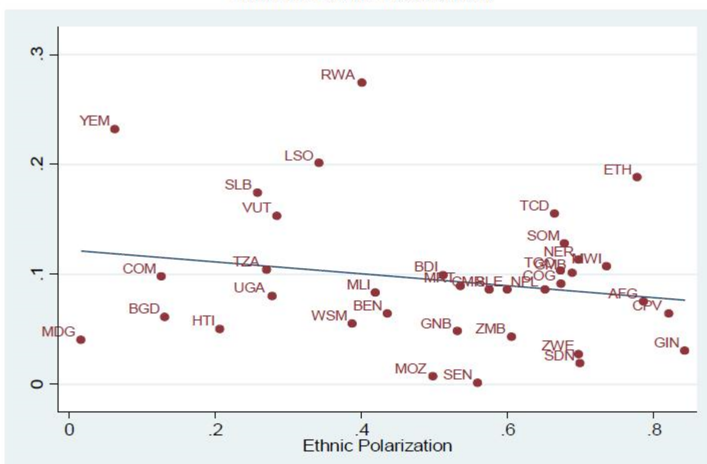
Figure 5: The Role of Ethnic Divisions

Panel A: Ethnic Fractionalization



Note: The figure shows the relationship between the country-specific slope estimates (reported in Figure 1) and ethnic fractionalization. The bootstrapped slope coefficient (s.e.) of the fitted regression line is -0.07 (0.03).

Panel B: Ethnic Polarization



Note: The figure shows the relationship between the country-specific slope estimates (reported in Figure 1) and ethnic polarization. The bootstrapped slope coefficient (s.e.) of the fitted regression line is -0.05 (0.05) .

Chapter 2: Fiscal Expansions Can Increase Unemployment (Joint with Evi Pappa)

1 Introduction

Most macroeconomists would agree that expansionary fiscal policy stimulates employment and lowers unemployment. Indeed, existing studies for the US economy (see, Ravn and Simonelli (2007) and Monacelli et al. (2010)) confirm this conventional wisdom. Our empirical analysis extends the literature by studying the effects of fiscal policy on unemployment in other OECD countries and shows, first, that increases in government spending can actually increase unemployment in many OECD countries and, second, that the existing evidence for the US is not robust to the sample period considered. We show that the unemployment response to government spending increases in the US is highly sensitive to the time period analyzed. For samples before 1980 and samples after 1995 unemployment decreases after a fiscal expansion, but for samples after 1980 the opposite is true. Our empirical work also shows that this time-heterogeneity is specific to the US. For the other OECD countries and all time periods we have considered increases in government spending are in general accompanied by increases in the unemployment rate. The fact that fiscal expansions increase unemployment is somewhat surprising. Yet, it is robust, in the sense that it holds for a number of OECD countries and sample periods and a variety of VAR specifications and identification schemes that one can use to extract fiscal shocks from the data.

Despite the difficulties in their identification, economists have tried to characterize the responses of macroeconomic variables such as investment, consumption and output to fiscal disturbances. Blanchard and Perotti (2002), Perotti (2004) and Gali et al. (2007) use the restriction that government spending does not contemporaneously react to changes in macrovariables to identify fiscal shocks. Ramey and Shapiro (1998), Edelberg et al. (1999), and Burnside et al. (2004) identify fiscal shocks as episodes of significant exogenous and unforeseen increases in government spending in national defense.[1] Canova and Pappa (2007) and Mountford and Uhlig (2009) identify fiscal shocks using sign restrictions. Pappa (2009a), using robust theoretical sign restrictions, was the first to investigate for the US the effects of fiscal shocks on labor market variables such as the real wage and employment. The analysis we conduct here considers many more labor market variables, covers as many as ten OECD countries, and focuses attention on the dynamics of the unemployment rate. Determining how the unemployment rate responds to fiscal expansions is important because many fiscal packages in the real world are typically designed to "create jobs" and because models have recently been proposed to explain its time series properties.

Our empirical analysis shows that the unemployment rate can increase sig-

[1] Depending on the identification approach the results on the effects of government spending on private consumption differ. Perotti (2007) critically reviews this literature.

nificantly in response to government expenditures shocks in many OECD countries. Results are robust to alternative identification schemes, the inclusion of control variables and different sub-periods for most countries but the US where the response of unemployment to fiscal shocks seems to have changed pattern substantially over time. In addition, we document that fiscal expansions tend to increase the participation rate, the employment rate and the real wage.

Our empirical findings are difficult to reconcile with existing theoretical models for several reasons. First, analyzing the effects of government spending shocks on unemployment in standard RBC and NK models is impossible since standard versions of these models only allow for movements in hours worked and/or employment. Second, even if we incorporate the Diamond-Mortensen-Pissarides search and matching model into standard frameworks, as suggested in Andolfatto (1996), or Walsh (2005), we cannot account for the responses of the participation rate – in these models participation is constant. But, even disregarding participation choices, simultaneously generating increases in output, real wages, the employment and the unemployment rate in response to fiscal shocks is difficult.

To circumvent these difficulties we add a participation margin in a New Keynesian model with labor market frictions as in Ravn (2008) and, in the spirit of Lindbeck and Snower (1988), we consider a labor market with insiders and outsiders. Endogenous participation generates an increase in the pool of job seekers after a fiscal expansion since the wealth effect induced by the shock in government´s absorption increases labor market participation. The assumptions on workers' heterogeneity and price stickiness are also crucial to generate increases in total employment and the unemployment rate. Sticky prices are necessary for inducing an increase in demand that counteracts the crowding out of vacancies due to the increase in government absorption. However, for low values of the labor supply elasticity participation does not increase enough and the increased labor demand by the sticky price firms is strong enough to fully absorb the supply of new participants. The fact that some new entrants, characterized as outsiders, have a less efficient matching technology guarantees that even for low values of the labor supply elasticity unemployment can increase.

Our paper is related to a number of recent works which have appeared in the literature. Relative to Monacelli et al. (2010), our model incorporates features such as endogenous participation and workers' heterogeneity that can generate increases in unemployment, output, employment and the real wage after a fiscal expansion. Faia et al. (2010) also assume that workers are heterogeneous and introduce labor frictions in the form of labor turnover costs but do not examine the dynamics of unemployment or labor participation in response to fiscal shocks. Finally, Gomes (2009) uses a two-sector dynamic stochastic general equilibrium model with search and matching frictions to study the labor market effects of shocks to public sector employment and wages. In his model unemployment decreases in response to generic government consumption shocks.

The remainder of the paper is organized as follows. Section 2 describes the econometric framework. Section 3 presents the main empirical results. The theoretical model is presented in Section 4. Section 5 describes the dynamics of the benchmark economy and highlights the features that are crucial for replicating qualitatively the empirical results and Section 6 concludes.

2 Data and Estimation Methodology

We obtain quarterly data on GDP, private consumption, private investment, government consumption expenditures, wages, the short-term interest rate, the labor force participation and the unemployment rate from OECD statistics. Total central government tax revenues are obtained for Canada from Statistics Canada, for Australia and Japan from Datastream, for the UK from the Office of National Statistics, and for the US from the Bureau of Economic Analysis. Except for the interest rate, the unemployment and the participation rate and real wages, all other variables are in real per capita terms and all variables are seasonally adjusted. The time periods cover the longest possible sample given by OECD statistics: (1968:1-2009:1) for Australia, (1961:1-2009:1) for Canada, (1990:1-2009:1) for Finland, (1978:1-2009:1) for France, (1980:1-2009:1) for Italy, (1980:1-2009:1) for Japan, (1979:1-2009:1) for Norway, (1982:1-2009:1) for Sweden, (1978:1-2009:1) for the UK, and (1964:1-2009:1) for the US. In the main part of the empirical analysis we focus on Australia, Canada, Japan, the UK, and the US because for these countries we have sufficiently long and reliable quarterly data on both government consumption expenditures and total government tax revenues.

To identify the impact that government expenditure shocks have on labor market outcomes we use a structural VAR approach. The variables entering our baseline specification are: the logs of real per capita government expenditures, GDP, consumption, and investment, and the interest rate, real CPI wage, and the unemployment rate. To start with we assume that government expenditures are contemporaneously unaffected by all variables in the model. This assumption appears plausible to us because fiscal policy usually reacts with at least a quarter lag to changes in the economic environment (see for instance Blanchard and Perotti, (2002); Perotti, (2004)). The lag length of our VAR model is based on information criteria and set equal to one. All variables in the VAR model enter as log-deviations from a constant and a quadratic time trend.[2] In all figures we report 95 percent confidence bands.

[2] We have checked the stability of our VAR by computing the eigenvalues of the estimated coefficient matrix. We found that all of the eigenvalues lie within the unit circle. We have also checked the robustness of our estimates using a VAR with up to 4 lags. Impulse responses from the 4-lag VAR are similar to our parsimonious 1-lag specification. Also, our results hold independently of the omission of the time trend in the specification. Responses for these specifications are available from the authors upon request.

3 Empirical Results

Panel A of Figure 1 presents unemployment responses for the baseline SVAR model. For all nine out of the ten OECD countries there is a significant positive response following increases in government expenditures. Government expenditure increases raise unemployment strongly in Finland and Sweden, while they induce no significant effects in Italy. The estimates imply that a 10% increase in government expenditures typically increases the unemployment rate at peak by around 0.2-0.5%. Responses are persistent, indicating that government expenditure increases may have effects on the unemployment rate that are of a long-lasting nature, which is in line with the hysteresis hypothesis (see Blanchard and Summers, (1987)).

In Panel B of Figure 1 we analyze what would happen if, rather than assuming that government expenditure is insensitive to economic conditions, we allow the government expenditure series to react to all VAR variables contemporaneously. Such an assumption can be justified by claiming that automatic stabilizers are present at any point in time. The responses displayed in the second panel of Figure 1 show that there continues to be a positive response in the unemployment rate following government expenditure increases: the increases are significant at the 95% confidence level for nine of the ten countries.

In the analysis so far we have not controlled for tax revenues in the VAR specification. This could be a crucial omission since it does not control for changes in the deficits and it does not rule out (potentially important) contemporaneous effects of distortionary tax changes on output. For that reason in what follows we focus the analysis on Australia, Canada, Japan, the UK, and the US. The variables entering our specification with tax revenues are: the logs of real per capita government expenditures, GDP, consumption, and investment, and the interest rate, real per capita tax revenues, real CPI wage, and the unemployment rate. Panel A of Figure 2 shows that, for all five OECD countries where we have data on tax revenues there is a significant positive response in the unemployment rate following expansionary government spending shocks. In Panel B of Figure 2 we analyze what would happen if, rather than assuming that government expenditure is insensitive to economic conditions, we allow the government expenditure series to react to all VAR variables contemporaneously. The responses displayed in the second panel of Figure 2 show that there continues to be a positive response in the unemployment rate following government expenditure increases at the 95% confidence level for Canada, Japan and the UK.

Given that different identification schemes might induce different dynamics in the endogenous variables, we have checked whether identifying fiscal shocks as unforeseen increases in government expenditure on defense, following the approach of Ramey and Shapiro (1998), changes the pattern of unemployment responses we obtained. In Figure 3 we return to our baseline VAR specification that includes tax revenues but substitute the government expenditure series for the Ramey-Shapiro war dummies for the US.[3] In this case, we also obtain a positive and statistically significant response in the unemployment rate that has its peak effect after about 2 quarters. The response in the US unemployment rate is persistent and turns negative only after about 10 quarters.

In contrast to our results, Ravn and Simonelli (2007) and Monacelli et al. (2010) find that, for the US, unemployment significantly decreases after an expansionary expenditure shock. It appears that the differences are due to the sample period used in the estimation. Ravn and Simonelli (2007) use data from 1959 to 2004, and Monacelli et al. (2010) use data from 1954 to 2006[4]. Perotti (2004) also finds that the effects of fiscal shocks change when considering the pre-80s and the post-80s samples. In order to investigate whether this is the case also for the unemployment response and in order to examine the robustness of our results to the subsample used we present in Figure 4 the unemployment responses for different subsample periods for Australia, Canada and the US (the three countries where we have long enough data to cover the pre-1980 period). Subsamples cover the periods 1968-1980, 1968-1985, 1968-1990, 1968-1995, 1968-2000, 1968-2005, and 1968-2009[5]. The behavior of unemployment responses to government expenditure increases is relatively unstable across subsamples for the US economy. Unemployment reacts negatively to expenditure increases up to the 1990s, and for longer subsamples the reaction is either insignificant or positive, while for Canada the responses are significantly positive regardless of the time-period covered and for Australia they are positive and significant in five of the seven subsamples considered.

To ensure that our results are not driven by a possible structural break that occurred around the turn of the 1980s, and that they are also robust to cross-country differences in the time period covered, we report in Figure 5 impulse responses for the periods 1980-2009, 1985-2009, 1990-2009, and 1995-2005. The responses of unemployment to government expenditure expansions in the US are unstable also for this time period. In the first three subsamples responses of unemployment to government spending shocks are significantly positive, while in the last period considered responses turn again negative and significant on impact. For the other OECD countries, with the exception of Japan where responses are almost never significant, unemployment increases following increases in government expenditures in almost all subsamples.

The increases in unemployment we document are accompanied by increases in output per capita: as the first panel of Figure 6 shows the impact response following the increase in government spending is positive in all countries. Thus, the increase in unemployment is not driven by a possibly adverse effect of the fiscal expansion on output. Interestingly, the estimated responses of private

[3]The Ramey-Shapiro war dummy takes on the value of 1 in 1965:1, 1980:1, 2001:3, and 2003:1. A first-stage regression of the change in government expenditures on the lagged Ramey-Shapiro war dummy yields during the 1964:1-2009:1 period a t-value of 2.73.

[4]In order to have reliable and comparable series we use OECD statistics data for all countries considered. However, our results for the US hold also for data from the BEA and the BLS, and for the data of Simonelli and Ravn (2007) available at: http://www.eui.eu/Personal/Ravn/.

[5]We start in 1968 to have the longest possible symmetric samples across the three countries.

consumption to these expenditure shocks (presented in Panel B of Figure 6) are also positive and significant for all countries except Japan. Instead, the shock crowds out private investment in all the countries but Australia (See Panel C of Figure 6).

Panel A of Figure 7 shows that increases in government spending also increase the real wage on impact. However, the responses are significant at the 95% confidence level only for the UK. In Panels B and C of Figure 7 we present the response of employment and the labor force participation rate.[6] Consistent with the findings in Pappa (2009b) we find that the employment rate significantly increases following a government consumption expenditure increase in the US and Canada. For Japan and the UK, employment rates significantly decrease while in Australia the response is insignificant. The response of the labor force participation rate is negative and significant for Japan, but positive and significant for the other four countries.

To strengthen our conclusions we have also checked whether results change when we identify fiscal shocks using sign restrictions on the responses of deficits, output, tax revenues and government expenditures. Following Pappa (2009b), we use the restriction that government expenditures, output and deficits are positively correlated contemporaneously, while tax revenues are not allowed to respond negatively to the shock.[7] In Figure 8 we plot the responses based on the sign restrictions identification. The unemployment rate significantly increases following government expenditure increases and responses are persistent for all five countries. Here real wages increase significantly after the fiscal expansion in all the countries and the responses of employment are insignificant in Japan and the UK. For the other countries results are very similar qualitatively with our baseline specification.

According to Gomes (2009) public sector wages may play an important role in shaping unemployment dynamics, since high public wages may induce unemployed to queue for public sector jobs. This is a relevant issue since a large component of government consumption expenditures corresponds to public wages. For example, public wages cover 52% of total government expenditures in Australia, 59% in Canada, 38% in Japan, 53% in the UK and 66.5% in the US. To exclude the possibility that unemployment increases are driven by increases in government wages, we have repeated our exercise replacing the government expenditure series with series of government consumption purchases. The results we obtain in Figure 9 are unchanged relative to the benchmark model.[8]

In economies where the expected present value of future taxes and expenditures matters for private sector agents' choices, current fiscal developments can have complex and sometimes surprising effects since current policy can play a crucial role in shaping expectations of future policy changes. So, for example,

[6]The impulses are generated from a VAR where we replace the unemployment rate by employment and the labor force (both variables are in per capita terms).

[7]Given that in the theoretical model we use output might not react contemporaneously to expenditure shocks we also use the above restrictions on the second period after the shock. Results are robust to this change as well.

[8]Australia is excluded due to unavailability of data on government consumption.

an expansionary fiscal shock may end up being contractionary if it induces sufficiently strong expectations of future policy changes in the opposite direction. To control for such effects we have repeated our exercise by including a forward looking variable like stock prices in the baseline VAR. As the first panel of Figure 10 shows, even when we control for changes in expectations, the effects of fiscal expansions on unemployment continue to be positive and significant for all countries (except for Japan). We have also made an attempt to further deal with anticipation effects by including changes of the international oil price in the VAR. Also these regressions produce a significant positive effect of government spending on the unemployment rate in all five countries (see Panel B of Figure 10).

To summarize: the evidence we have collected indicates that fiscal expansions can increase the real wage, the employment and the labor force participation rate together with output and the unemployment rate. This evidence is hard to reconcile with standard models. The fact that increases in government spending increase unemployment and the labor force participation rate gives us a starting point to search for potentially consistent theoretical explanations. In the next section we describe how a model with endogenous labor force participation and insiders and outsiders can account for these facts.

4 The Model

Analyzing the effects of government spending shocks on unemployment, or the participation rate in standard models is hard since most models allow only for voluntary movements in hours of work and employment. To analyze unemployment fluctuations researchers found it natural to incorporate the Diamond (1982) and Mortensen and Pissarides (1994) search and matching model into the standard frameworks. Among others, Andolfatto (1996), den Haan, Ramey and Watson (2000), Shimer (2005) and Ravn (2008) have introduced search frictions into a standard RBC model. Walsh (2005), Trigari (2009), Campolmi and Faia (forthcoming), Thomas (2008) and Blanchard and Gali (2010) have added them to New Keynesian models.

However, these studies assume that the labor market participation rate is constant. The empirical analysis has revealed that government spending shocks do affect labor force participation. Hence, it is central to introduce a participation margin in our theoretical model. Following Ravn (2008), we model the labor market participation choice in terms of a trade-off between the reduction in leisure time to participate in the labor market search and the benefits associated with the prospect of finding a new job. Labor market non-participants are modeled as agents that are unmatched and that do not currently look for a job, while unemployed are unmatched agents that actively look for a job.

The traditional macroeconomic literature on unemployment (see Layard et al. (1991) for a literature review) discusses many reasons for why unemployment may occur in equilibrium. Lindbeck and Snower (1988) propose a model of insiders and outsiders for explaining unemployment. In their framework, unemployment occurs because some agents (the outsiders) cannot sell as much labor services as they wish to supply. We find this set up attractive, since in the real world many classes of agents, such as long-term unemployed, spouses, students, or elderly workers may be viewed as outsiders in the sense of Lindbeck and Snower (1988). These agents may often decide not to participate in the labor market and they might differ from the typical unemployed worker in their matching market prospects. Thus, the expected payoff from engaging in search activities is smaller for labor market non-participants (outsiders) than for search active agents (insiders). To incorporate the notion of insiders and outsiders in our model we introduce heterogeneity in the matching function. In particular, we assume that there are two types of unemployed workers that differ in their prospect of being matched with vacancies, with outsiders facing a less efficient matching technology than insiders. Finally, we will assume that prices are sticky in the short run, as a short-cut for generating a demand effect after a government spending shock.

The economy consists of households that have employed, unemployed and non-participants members. There are two types of firms in the economy: (i) competitive intermediate firms that use capital and labor to produce a good, and (ii) monopolistic competitive retailers that use all intermediate varieties to produce the final good which is then used for consumption, investment and government spending. Price rigidities arise at the retail level, while search frictions occur in the intermediate goods sector.

4.1 Preferences

There is a measure one of households. Households consist of a continuum of agents and the number of individuals in the household is large enough to guarantee insurance over consumption of its members.

At any point in time a fraction n_t of the household´s members are employed, a fraction u_t are unemployed and a fraction l_t are labor market non-participants. The difference between non-participants and unemployed is that the latter are actively looking for a job.

$$1 = n_t + u_t + l_t \tag{1}$$

The preferences of the representative household are defined by:

$$u(c_t, l_t) = \frac{c_t^{1-\eta}}{1-\eta} + \Phi \frac{l_t^{1-\zeta}}{1-\zeta} \tag{2}$$

where c_t, denotes consumption, $1/\eta$ is the intertemporal elasticity of substitution, $\Phi > 0$ is a preference parameter and ζ is the inverse of the elasticity of labor supply. That is, households obtain utility from consumption and from the fraction of households that do not participate in market activities and enjoy leisure.[9] Notice that each household member´s consumption is the same inde-

[9]Such a utility function can be rationalized by the production of home goods. That is, it is equivalent to assuming that households derive utility from market and home goods, c_t^h whereas the home goods are produced by the following production function: $c_t^h = \frac{l_t^{1-\zeta}}{1-\zeta}$.

pendently of their labor market status due to income pooling. Notice also that a member of a household that searches for a job or that is employed suffers the same disutility. That is, search effort is as costly in terms of utility as a full time job.

4.2 Matching

The process through which workers and firms find each other is represented by a matching function that accounts for the imperfections and transaction costs in the labor market.

We model heterogeneity in the matching functions of insiders and outsiders as follows. Every period a constant fraction σ of the currently employed worker-job matches is destroyed and a measure of M new matches are formed. Workers that experience a termination of their match are characterized as insiders and they enter into a period of unemployment. An insider may either remain unemployed, find a new job match, or become an outsider. Insiders become outsiders with probability $\mu \in [0,1]$. The number of new matches between vacant jobs and unmatched agents will depend on both the labor market tightness and the structure of unemployment. The aggregate number of matches is given by:

$$\begin{aligned} M(v_t, u_t^O, u_t^I) &= m_I(v_t, u_t^I) + m_O(v_t, u_t^O), \quad \text{with} \\ m_I(v,u) &> m_O(v,u) \;\; \text{for} \;\; \forall\, v, u > 0 \end{aligned} \tag{3}$$

where v denotes vacancies, u^I denotes the measure of insiders, while u^O denotes the measure of outsiders looking for a job. We assume that the efficiency of the matching process is higher for unemployed insiders than for unemployed outsiders. Thus, the matching function for the two groups of individuals is assumed to satisfy:

$$m_j(v, u^j) = \varrho_m^j v^\alpha (u^j)^{1-\alpha} \text{ with } \; j = I, O \quad \text{and} \varrho_m^I > \varrho_m^O > 0 \tag{4}$$

The probability that a vacant job is matched with a worker is going to depend on the overall labor market tightness, $\theta_t = \frac{v_t}{u_t}$, as in the standard framework, and on the relative size of insiders and outsiders. If we denote by γ_t^f this probability, we have:

$$\gamma_t^f = \frac{m_t}{v_t} = \theta_t^{\alpha-1}\left[\varrho_m^I\left(\frac{u_t^I}{u_t}\right)^{1-\alpha} + \varrho_m^O\left(\frac{u_t^O}{u_t}\right)^{1-\alpha}\right] \tag{5}$$

where $u = u_I + u_O$, and the ratio $\frac{u^j}{u}$, $j = I, O$, defines the share of unemployment for agents of type i. Thus, an increase in the unemployment rate for each type of agents increases the probability that a vacancy will be filled. However, an increase in the unemployment rate for insiders has a stronger impact on this probability than an increase in the unemployment rate of outsiders. The probability for an unemployed worker (insider or outsider) to find a job is:

$$\gamma_t^h = \frac{m_t}{u_t} = \theta_t^\alpha \left[\varrho_m^I \left(\frac{u^I}{u} \right)^{1-\alpha} + \varrho_m^O \left(\frac{u^O}{u} \right)^{1-\alpha} \right] \tag{6}$$

Again, the relative size of the two types of unemployed workers in the economy matters. Hence, an additional outsider searcher creates less of a negative externality for the total sum of individuals looking for a job. The probabilities to find a job for each type of agents are given by:

$$\gamma_{ji}^h = \frac{m_{jt}}{u_t}, j = O, I \tag{7}$$

The employment transition equation is given by:

$$n_{t+1} = (1-\sigma)n_t + m_{It} + m_{Ot} \tag{8}$$

The transition equation for insiders' unemployment is given by:

$$u_{t+1}^I = (1-\mu)u_t^I + \sigma n_t - m_{It} \tag{9}$$

Notice that insiders are more often (that is, for many parameter specifications) better off searching than non-participating since they are faced with a better matching technology. Outsiders instead have to decide whether they should participate in the labor market and their decision takes into account the fact that they are less advantageous in matching with firms.

4.3 The problem of the household

The household owns the economy´s capital stock. The capital stock evolves over time according to:

$$k_{t+1} = (1-\delta)k_t + i_t + \xi(\frac{k_{t+1}}{k_t})k_{t+1} \tag{10}$$

where δ is the capital´s depreciation rate, i_t is gross investment and $\xi(.)$ is a function that regulates capital adjustment costs. We adopt a quadratic specification of the form:

$$\xi(\frac{k_{t+1}}{k_t}) = \frac{\omega}{2}\left(\frac{k_{t+1}}{k_t} - 1 \right)^2 \tag{11}$$

where the parameter ω regulates the importance of capital adjustment costs for the accumulation of capital.

The representative household maximizes its expected utility given by:

$$E_t \sum_{t=0}^{\infty} \beta^t u(c_t, l_t) \tag{12}$$

choosing sequences of consumption, c_t, the number of insiders in the next period, u_{t+1}^I, and the number of outsiders, u_t^O, employment for next period,

n_{t+1}, next period´s bond holdings, B_{t+1} and capital, k_{t+1}, subject to (1), (8), (9), (10) and its budget constraint given by:

$$c_t + i_t + \frac{B_{t+1}}{p_t R_t} \leq r_t k_t + w_t n_t + b u_t + \frac{B_t}{p_t} + \Pi_t - T_t \tag{13}$$

where p_t is the price level, w_t is the real wage, r_t is the real return to capital, b denotes some non-tradable value to being unemployed expressed in terms of unit output, R_t is the gross nominal interest rate, Π_t are the profits of the monopolistic competitive firms and T_t are lump sum taxes paid to the government.

4.4 Intermediate goods firms and job creation

Intermediate goods firms employ the household's labor and capital to produce intermediate goods. The production function for intermediate goods is given by:

$$y_t = F(k_t, n_t) = k_t^{\varphi} {n_t}^{1-\varphi} \tag{14}$$

Intermediate firms maximize the discounted value of future profits. Firms adjust employment by varying the number of workers (extensive margin) rather than the number of hours per worker. According to Hansen (1985), most of the employment fluctuations arise from movements in this margin. The firm takes as given the number of workers currently employed and its employment decision concerns the number of vacancies that it posts in the current period, v_t. Firms open as many vacancies as necessary to employ the desired number of workers next period and there is a utility cost from posting a vacancy, $\varkappa$. Firms also need to decide on the size of the capital stock that they need for production. The problem of a firm with n_t currently employed workers consists of choosing capital and vacancies to maximize:

$$Q(n_t, k_t) = \max x_t F(k_t, n_t) - w_t n_t - r_t k_t - \varkappa v_t + E_t \Lambda_{t+1} Q(n_{t+1}, k_{t+1}) \tag{15}$$

where x_t is the relative price of intermediate goods and $\Lambda_{t+s} = \frac{\beta^s U_{ct+s}}{U_{ct}}$, is the discount factor. The maximization takes place subject to the production function, the law of motion for aggregate productivity and the job transition function that links the future number of filled jobs to the current stock of filled jobs plus net hiring.

$$n_{t+1} = (1-\sigma) n_t + \gamma_t^f v_t \tag{16}$$

4.5 Bargaining over wages

Workers and firms split rents through Nash bargaining and the part of the surplus they receive depends on their bargaining power. If we denote by $\vartheta \in (0,1)$ the firms bargaining power, the Nash bargaining problem is to maximize the weighted sum of log surpluses:

$$\max_{w_t} (1-\vartheta) \ln V_t^W + \vartheta \ln V_t^F$$

where $V_t^W = w_t - b + (1-\sigma-(\psi_t^{Ih}+\psi_t^{Oh}))E_t\Lambda_{t+1}V_{t+1}^W$, is the worker's surplus and $V_t^F = x_t(1-\varphi)\frac{y_t}{n_t} - w_t + \beta E_t\Lambda_{t+1}V_{t+1}^F$, is the firm's surplus of the match.

The solution of the bargaining problem defines the contractual wage as:

$$w_t = (1-\vartheta)\left[(1-\varphi)x_t\frac{y_t}{n_t} + \frac{\varkappa(\psi_t^{Oh}+\psi_t^{Ih})}{\gamma_t^f}\right] + \vartheta b \tag{17}$$

Note that in equilibrium, the value of working is the same for insiders and outsiders because otherwise firms could make profits by hiring less of those workers with a lower value and more of those workers with a higher value. In other words, there are decreasing returns in matching to unemployment, so in equilibrium the value of work should be the same in order for there to be no arbitrage opportunities. The wage paid to matched unemployed insiders will therefore be the same as the wage paid to matched unemployed outsiders.

4.6 Retailers and price setting

There is a continuum of monopolistically competitive retailers indexed by i on the unit interval. Retailers buy intermediate goods from firms and differentiate them with a technology that transforms one unit of intermediate goods into one unit of retail goods. Retail goods are then used for consumption, government spending and investment. Note that the relative price of intermediate goods, x_t, coincides with the real marginal cost faced by the retailers. Let y_{it} be the quantity of output sold by retailer i. Final goods can be expressed as the composite of individual retail goods:

$$y_t = \left[\int_0^1 y_{it}^{\frac{\varepsilon-1}{\varepsilon}} di\right]^{\frac{\varepsilon}{\varepsilon-1}} \tag{18}$$

where $\varepsilon > 1$ is the constant elasticity of demand for intermediate goods. The retail good is sold at its price, $p_t = \left(\int_0^1 p_{it}^{1-\varepsilon} di\right)^{\frac{1}{1-\varepsilon}}$. The resulting demand for each intermediate good depends on its relative price and aggregate demand:

$$y_{it} = \left(\frac{p_{it}}{p_t}\right)^{-\varepsilon} y_t \tag{19}$$

Following Calvo (1983) we assume that in any given period each retailer can reset its price with a fixed probability $1-\chi$. Hence, the price index is given by:

$$p_t = \left[(1-\chi)p_t^{*1-\varepsilon} + \chi p_{t-1}^{1-\varepsilon}\right]^{1/(1-\varepsilon)} \tag{20}$$

The firms that are able to reset their price, p_t^*, choose it so as to maximize expected profits given by:

$$E_t\sum_{t=0}^{\infty}\chi^s\Lambda_{t+s}\left[\frac{p_{it}^*}{p_{t+s}} - x_{t+s}\right]y_{it+s} \tag{21}$$

4.7 Fiscal policy

The government consumes exogenously part of the retail goods and finances its expenditures via lump sum taxes.

$$bu_t + G_t = T_t$$

4.8 Monetary Policy

There is an independent monetary authority which sets the nominal interest rate as a function of current inflation, according to the rule:

$$R_t = \overline{R}\exp(\zeta_\pi \pi_t) \tag{22}$$

where π_t measures inflation in deviation from the steady state.

4.9 Closing the model

Aggregate production must equal private and public demand:

$$y_t = c_t + i_t + G_t + \varkappa v_t \tag{23}$$

4.10 Parameterization

We solve the model by approximating the equilibrium conditions around a non-stochastic steady state in which all prices are flexible. The full list of our parameter choices is given in Table 1. The quarterly discount factor is set to 0.99, which implies a quarterly real rate of interest of approximately 1 percent. The risk aversion parameter η is set to 2 and the utility of leisure has elasticity $\zeta = 4$. The implied value of the labor supply is somewhat lower than what researchers usually assume in the literature. In the next section we will show that workers' heterogeneity is key for using low values of this elasticity.

Following Blanchard and Diamond (1989) we set $\alpha = 0.6$ and, using Hosios condition, we also set the bargaining parameter equal to the elasticity of matching, i.e., $\alpha = \vartheta$.

Davis, Haltiwanger and Schuh (1996) compute a quarterly worker separation rate of about 8 percent, while Hall (1995) reports this rate to be between 8 and 10 percent. Thus, we set the separation rate parameter σ to 0.09. The probability of becoming an outsider is set to 0.1. With this parameterization we match that long term unemployment (outsiders) represents 21% of total unemployment, in line with CPS data. The values of ρ_m^O and ρ_m^I are set so that the total unemployment rate and the market tightness equal 7% and 0.25, respectively. The level of benefits in the steady state is set so that labor force participation equals 70%; the vacancy to output ratio is set equal to 0.01.

The depreciation rate is set equal to 0.025 and the capital share is set equal to 0.36. Capital adjustment costs are included to moderate the response of investment with respect to fiscal shocks. We set parameter ω to match the ratio of the investment to output variance for the US economy when we include TFP and monetary shocks in the model. The probability that a firm does not change its price within a given period, ψ, is set equal to 0.75, implying that the average period between price adjustments is around 4 quarters. The value used for the persistence of the government spending shock is the average of the cross country values we have obtained in Section 3.

5 How expansionary government spending shocks increase unemployment

We first investigate the properties of the benchmark model and examine the mechanisms leading to the results of interest.

5.1 The benchmark model

Figure 11 presents the effects of a government expenditure shock on output, employment, unemployment (total and for the two types of workers), the real wage, the participation rate, consumption and investment.

An increase in government spending induces a negative wealth effect that makes households increase their labor supply. As a result, the participation rate increases. Also, the increase in government absorption is crowding out private consumption, investment and hiring. On the other hand, the increase in demand induced by the government expansion increases labor demand, and, in turn, wages and employment increase. Non-participants evaluate that it is good to invest in search when government spending increases since there is the extra benefit of facing the more efficient search technology after an employment spell. But, since it is the insiders that get the extra jobs, the unemployment rate of the outsiders increases. Consequently, total unemployment increases on impact because of the increase in participation and the increase in the unemployment rate of outsiders. As insiders are hired by the firms to face the increased demand, total unemployment decreases; but when the demand effect fades away total unemployment starts rising again. In line with the empirical results, the responses of unemployment are very persistent.

5.2 The role of price stickiness

Price stickiness is necessary for obtaining our results. In Figure 12 we present the responses of an economy which is otherwise identical to the benchmark except for the assumption of price stickiness. With flexible prices, the increase in government absorption would crowd out vacancy posting (as it crowds out consumption and investment) since it would decrease the resources available for filling vacancies. Although the wealth effect of the shock would increase participation and the labor supply in equilibrium, the decrease in vacancy posting would decrease demand for employment and output and increase the unemployment of both types of agents, generating output and employment responses which are in contrast with the empirical evidence we have reported.

5.3 The role of the participation margin and workers' heterogeneity

We have modeled the participation margin in order to be able to analyze the behavior of labor force participation in reaction to expenditure shocks. However, the use of the participation margin might be important in generating the results. In Figure 13 we plot the responses of the variables when agents are homogeneous, prices are sticky and there is no participation margin and when agents are homogeneous, prices are sticky and there is a participation margin.[10] The fact that there is a pool of non-participants that move into the labor force when the negative wealth effect from the increase in the government absorption kicks in is not enough to generate an increase in unemployment after a government spending shock. In fact, the two models with or without the participation decision would be almost identical. Workers' heterogeneity is crucial for generating the increase in total unemployment after the spending shock for low values of the labor supply elasticity. If agents were homogeneous, an increase in government spending would increase labor demand and unemployment would be reduced. It is the fact that outsiders have a hard time to find a job relative to the insiders that makes total unemployment increase in equilibrium when the labor supply elasticity is low.

For higher values of the labor supply elasticity, the presence of a participation margin would be sufficient to generate increases in unemployment after a fiscal expansion. We show this in Figure 14 where we plot the response of unemployment in the homogeneous agents model when we vary ζ, the variable determining the Frisch elasticity, $1/\zeta$. For high values of the labor supply elasticity, the wealth effect increases participation and makes unemployment increase even when agents are homogenous. Thus, while both the presence of the labor participation margin and workers' heterogeneity matter, the latter is crucial for generating a positive response of total unemployment when the labor supply elasticity takes low values.

5.4 Other important features

We performed a number of sensitivity analysis exercises to investigate the robustness of our conclusions with respect to changes in the remaining parameters

[10] For the homogeneous workers' model, the variable u^I disappears and $u^O = u$. The matching function is given by: $m_t = \rho_m v^\alpha u^{1-\alpha}$ and agents maximize:

$$u(c_t, n_t) = \frac{c_t^{1-\eta}}{1-\eta} - \Phi \frac{n_t^{1-\zeta}}{1-\zeta}$$

subject to (10), (8), and (13), and (1) becomes: $n_t + u_t = 1$.

With the participation margin, agents solve the same problem as in the benchmark economy with the only difference that $u^I = 0$. All models are parameterized to deliver comparable steady state values for the labor market variables.

of the model. The most crucial parameters for the dynamics of unemployment are the cost of posting a vacancy as a percentage of GDP, $\varkappa$, the adjustment cost parameter, ω, the labor supply elasticity, $1/\zeta$ and the relative size of outsiders to total unemployment.

The size of the vacancy cost is important to determine how much the government expansion crowds out the creation of vacancies. If the cost associated with the creation of vacancies is very small ($\varkappa = 0.001$ in Panel A of Figure 14) an increase in government spending does not crowd out substantially job creation and the wealth and the demand effects lead to increases in employment and vacancies, decreasing unemployment for both types of workers.

The presence of capital adjustment costs ensures that the crowding out of investment is limited so that capital and employment do not fall after the expenditure expansion. The size of capital adjustment costs affects the magnitude of the initial response to the shock as well as its persistence since it affects the accumulation of capital. The sensitivity of total unemployment responses to changes in ω is presented in Panel B of Figure 15. Notice that in the model with no capital ($\omega \to \infty$) the wealth effect of the increase in government absorption becomes stronger and unemployment increases significantly on impact after the fiscal expansion.

On the other hand, when the labor supply elasticity decreases (for values of $\zeta \geq 10$), the wealth effect of the increase in government absorption does not increase labor force participation significantly. As a result, the unemployed of both types can be employed in firms that face increased demand for their products and unemployment decreases instantaneously after the fiscal expansion (see Panel C of Figure 15).

Finally the relative size of insiders and outsiders in total unemployment matters. The relative size of outsiders and insiders in total unemployment is determined by the parameter μ. Panel D of Figure 15 plots the responses of total unemployment when we vary μ. Unemployment decreases after an expansionary expenditure shock for $0.1 < \mu < 0.95$, or, in other words, if the share of outsiders in total unemployment varies between [9%, 90%]. When the share of outsiders in total unemployment is below 9%, expenditure increases lead to a fall in total unemployment.

Hence, for low values of the relative size of outsiders, the model predicts reductions in total unemployment after a fiscal expansion. Interestingly for the US the share of long-term unemployed to total unemployed is significantly different for the different subsamples considered. According to the Labor Force Statistics from the Current Population Survey, the average value of the percentage of unemployed with unemployment duration higher than 27 weeks in total unemployment is 14.3% for the sample period 1954:2010, while for the sample period 1954-1979 it is 10% and for the period 1980:2010 it accounts for 17.3% of total unemployment. Moreover, in the beginning of the 1980s this percentage equals 20% and is pretty much higher relative to its average value. This evidence squares well with our theoretical model since it can explain the changes in the behavior of unemployment in response to fiscal shocks over time in the US by changes in the share of outsiders (viewed as long-term unemployed) in this country.

5.5 The response of private consumption

Our model was designed to show that it is possible to generate an increase in total unemployment after an expenditure expansion under reasonable assumptions and the goal of the previous section has been to highlight the elements needed to reproduce the empirical regularities. However, the proposed model fails to account for the consumption dynamics we have presented in Section 3. In particular, in the model private consumption decreases after expenditure increases. Given that we are primarily concerned with reproducing the dynamics of the labor market after an expenditure increase, we have not included in the baseline model mechanisms that would overcome this shortcoming. Here we show that if government and private consumption are complements as in Linnemann and Schaubert (2003), the theoretical consumption responses become more consistent with the data.[11]

Preferences are now defined by:

$$u(c_t, l_t) = \frac{\left[\left\{\nu c_t^{\frac{\xi-1}{\xi}} + (1-\nu)G_t^{\frac{\xi-1}{\xi}}\right\}^{\frac{\xi}{\xi-1}}\right]^{1-\eta}}{1-\eta} + \Phi\frac{l_t^{1-\zeta}}{1-\zeta}$$

where the degree of substitutability between private and public consumption is regulated by ξ. The share parameter ν determines how much public consumption affects utility: when $\nu = 1$, public consumption is useless from the agents' point of view and the model is identical to the baseline specification.

In Figure 16 we present responses when ν=0.7 and $\xi = 0.4$. When public and private consumption are complements an increase in government expenditures increases private consumption at the expense of a larger crowding out of investment in equilibrium. At the same time, the complementarity between private and public consumption does not cancel out the negative wealth effect due to the increase in government's absorption and labor force participation increases generating an increase in total unemployment in equilibrium.

6 Conclusions

We empirically examined the effect of government expenditure shocks on labor market variables and, in particular, on unemployment for OECD countries and found that a fiscal expansion can lead to a significant increase in the unemployment rate for many countries and many of the time periods considered. We have

[11] Complementarity between consumption and leisure (see Hall and Milgrom (2005)) could in principle generate increases in private consumption after a fiscal shock. Shimer (2010) shows how to incorporate income pooling with non separable utility between consumption and leisure. However, when a labor participation margin is allowed the utility specification used by Shimer is not easily applicable.

shown that results are robust to the identification scheme used to extract fiscal shocks from the data, the subsample period, and the inclusion of additional control variables for all countries that we have available data, except for the US. The responses of unemployment to government spending shocks is sensitive to the time period analyzed.

Our empirical results suggest, against the common wisdom, that government expansions can lead to increases in unemployment. Following a recent trend we consider a New Keynesian model with search frictions, endogenous participation and workers' heterogeneity to explain the empirical findings. In contrast to the existing literature, our model can generate depending on the exact parametrization, positive or negative responses of unemployment in response to positive government spending shocks and can possibly explain the reason behind the differences in the unemployment responses to government spending shocks in the US subsamples. The introduction of workers' heterogeneity is crucial for deriving our results. When the economy is populated by insiders and outsiders facing different matching prospects in the labor market, total unemployment may increase after a fiscal expansion. This is because the negative wealth effect induced by the increase in government absorption increases labor force participation. However, outsiders unemployment increases more than the fall in insiders unemployment and total unemployment increases in equilibrium.

While our empirical analysis is potentially subject to the standard critiques raised to VAR exercises (see, e.g., Chari et al. (2007) and Ramey (2009)) it is unlikely that empirical analysis conducted with different tools will lead to results that are different from those we have since the dynamics of unemployment we present are robust to different identification schemes, possible controls for anticipated effects and specifications of the VAR. Thus, any model with features different from those we consider must be compared with the particular stylized facts we present.

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8 Appendix

8.1 First order conditions

Households

The first order conditions for the household are given below:

$$c_t^{-\eta} = \lambda_{ct}$$

$$\lambda_{ct}\left(1+\omega\left[\frac{k_{t+1}}{k_t}-1\right]\right) = \beta E_t \lambda_{ct+1}\left(1-\delta+r_{t+1}+\frac{\omega}{2}\left(\frac{k_{t+2}}{k_{t+1}}-1\right)^2\right)$$

$$\phi l_t^{-\zeta} = \psi_t^{Oh}\lambda_{nt} + b\lambda_{ct}$$

$$\lambda_{ut} = \beta E_t[\lambda_{nt+1}\psi_{t+1}^{Ih} + \lambda_{ct+1}b + \lambda_{ut+1}((1-\mu)-\psi_{t+1}^{Ih}) - \phi l_{t+1}^{-\zeta}]$$

$$\lambda_{nt} = \beta E_t[\lambda_{ct+1}w_{t+1} + (1-\sigma)\lambda_{nt+1} + \sigma\lambda_{ut+1} - \phi l_{t+1}^{-\zeta}]$$

$$\lambda_{ct}\pi_{t+1} = \beta E_t \lambda_{ct+1} R_t$$

where $\psi_t^{Ih} = \frac{m_t^I}{u_t^I}$ and $\psi_t^{Oh} = \frac{m_t^O}{u_t^O}$.

Intermediate firms

The first order conditions for the firm are given by:

$$x_t F_{kt} = r_t$$

$$\frac{\varkappa}{\gamma_t^f} = \beta E_t\left(\frac{c_t}{c_{t+1}}\right)^\eta\left[x_{t+1}F_{nt+1} - w_{t+1} + (1-\sigma)\frac{\varkappa}{\gamma_{t+1}^f}\right]$$

Retailers

The optimal price for a retailer that can reset her price in the current period solves:

$$p_{it}^* = \frac{\varepsilon}{\varepsilon-1}\frac{E_t\sum_{t=0}^{\infty}\chi^s\Lambda_{t+s}x_{t+s}y_{it+s}}{E_t\sum_{t=0}^{\infty}\chi^s\Lambda_{t+s}y_{it+s}} \tag{24}$$

8.2 Steady state

The steady state is one with no employment, or unemployment growth and zero inflation.

$$\sigma n = \psi^{Ih}u_I + \psi^{Oh}u_O \tag{25}$$

$$\mu u_I = \psi^{Oh}u_O \tag{26}$$

$$1/\beta = 1 - \delta + r \tag{27}$$

$$1/\beta = R \tag{28}$$

$$\phi l^{-\zeta} = \psi^{Oh}\lambda_n + bc^{-\eta} \tag{29}$$

$$\lambda_u = \beta[\lambda_n \psi^{Ih} + \lambda_u((1-\mu) - \psi^{Ih}) - \phi l^{-\zeta} + bc^{-\eta}] \tag{30}$$

$$\lambda_n = \beta[c^{-\eta}w + (1-\sigma)\lambda_n + \sigma\lambda_u - \phi l^{-\zeta}] \tag{31}$$

$$r = \varphi x \frac{y}{k} \tag{32}$$

$$\frac{i}{k} = \delta \tag{33}$$

$$y = Zk^{\varphi}(n)^{1-\varphi}, Z = 1 \tag{34}$$

$$1 = l + u_I + u_O + n \tag{35}$$

$$y = c + i + g + \varkappa v \tag{36}$$

$$\theta_O = v/u_O, \qquad \theta_I = v/u_I \tag{37}$$

$$\psi^{Ih} = \rho_m^I \theta_I^\alpha, \qquad \psi^{Oh} = \rho_m^O \theta_O^\alpha$$

$$\begin{aligned} \gamma^f &= \theta^{\alpha-1}\left[\varrho_m^I\left(\frac{u^I}{u}\right)^{1-\alpha} + \varrho_m^O\left(\frac{u^O}{u}\right)^{1-\alpha}\right] && (38) \\ \gamma^f &= \psi^{If} + \psi^{Of} \text{ with} && (39) \\ \psi^{If} &= \psi^{Ih}/\theta_I \psi^{Of} = \psi^{Oh}/\theta_O && (40) \end{aligned}$$

$$\frac{\varkappa}{\gamma^f}(1 - \beta(1-\sigma)) = \beta\left[\frac{y}{n}x(1-\varphi) - w\right] \tag{41}$$

$$w = (1-\vartheta)\left[(1-\varphi)x\frac{y}{n} + \frac{\varkappa(\psi^{Oh} + \psi^{Ih})}{\gamma^f}\right] + \vartheta b \tag{42}$$

$$x = \frac{\varepsilon - 1}{\varepsilon} \tag{43}$$

Substituting (25) and (26) and the fact that $\frac{\theta_I}{\theta_O} = \frac{u_O}{u_I}$ in the remaining equations we get:

$$n = u_O \frac{\rho_m^O \theta_O^\alpha}{\mu\sigma}[\rho_m^I \left(\frac{\mu}{\rho_m^O}\right)^\alpha \theta_O^{\alpha(1-\alpha)} + \mu] = B(\theta_O)u_O \tag{44}$$

$$\lambda_n = \frac{c^{-\eta}(w-b)}{\frac{1}{\beta} - (1-\sigma) + \rho_m^O \theta_O^\alpha - \frac{\beta\sigma(\rho_m^I \theta_I^\alpha - \rho_m^O \theta_O^\alpha)}{1-(1-\mu)\beta + \beta\rho_m^I \theta_I^\alpha}} = \frac{c^{-\eta}(w-b)}{T(\theta_O)} \tag{45}$$

$$\theta_I = \left(\frac{\mu}{\rho_m^O}\right)\theta_O^{(1-\alpha)} \qquad \text{that is } \theta_I(\theta_O) \tag{46}$$

$$\lambda_u = \lambda_n \frac{\beta(\rho_m^I \theta_I^\alpha - \rho_m^O \theta_O^\alpha)}{1-\beta(1-\mu) - \beta\rho_m^I \theta_I^\alpha} \qquad \text{that is } \lambda_u(\theta_O) \tag{47}$$

$$\frac{y}{n} = \left[\frac{y}{k}\right]^{\frac{\varphi}{1-\varphi}} = \left[\frac{r}{\varphi}\frac{\varepsilon}{\varepsilon-1}\right]^{\frac{\varphi}{1-\varphi}} \tag{48}$$

from (35) we have

$$l = 1 - \left[1 + B(\theta_O) + \frac{\psi^{Oh}}{\mu}\right] u_O \tag{49}$$

We can write the resource constraint as:

$$\frac{c}{y} = 1 - \frac{\delta}{\frac{y}{k}} + \frac{g}{y} - \frac{\varkappa}{y}\theta_O u_O \tag{50}$$

and $c = \frac{c}{y}y$, while $y = \frac{y}{n}n$.

from (45) we have:

$$w = c^\eta \lambda_n T(\theta_O) + b \tag{51}$$

Using (41) together with (42) we can write:

$$u_O = \frac{\beta\vartheta\left[1-\varphi)x\frac{y}{n} - b\right]}{1-\beta(1-\sigma) - (1-\alpha)\beta(\psi^{Oh} + \psi^{Ih})} \frac{\gamma^f}{\frac{y}{n}T(\theta_O)\frac{\varkappa}{y}}$$

then using the equation for wages:

$$w = (1-\vartheta)\left[(1-\varphi)x\frac{y}{n} + \frac{\varkappa(\psi^{Oh} + \psi^{Ih})}{\gamma^f}\right] + \vartheta b$$

and equation (51) we have one equation in one unknown θ_O and its solution solves for the steady state of the model.

8.3 Loglinear conditions

State variables are 3: capital, employment and insider unemployment.

$$\widehat{n}_{t+1} = (1-\sigma)\widehat{n}_t + \frac{m_I}{n}\widehat{m}_{It} + \frac{m_O}{n}\widehat{m}_{Ot} \tag{A1}$$

$$\widehat{m}_{It} = \alpha\widehat{v}_t + (1-\alpha)\widehat{u}_t^I \tag{A2}$$

$$\widehat{m}_{Ot} = \alpha\widehat{v}_t + (1-\alpha)\widehat{u}_t^O \tag{A3}$$

$$\widehat{\psi}_t^{Ih} = \widehat{m}_{It} - \widehat{u}_t^I \tag{A4}$$

$$\widehat{\psi}_t^{Oh} = \widehat{m}_{Ot} - \widehat{u}_t^O \tag{A5}$$

$$\widehat{k}_{t+1} = (1-\delta)\widehat{k}_t + \delta\widehat{i}_t \tag{A6}$$

$$\widehat{u}_{t+1}^I = (1-\mu)\widehat{u}_t^I + \sigma\frac{n}{u^I}\widehat{n}_t - \frac{m_I}{u^I}\widehat{m}_{It} \tag{A7}$$

$$l\widehat{l}_t + n\widehat{n}_t + u^I\widehat{u}_t^I + u^O\widehat{u}_t^0 = 0 \tag{A8}$$

$$\frac{\eta}{\beta}\widehat{c}_t + \frac{\omega}{\beta}k_t = E_t\{\frac{\eta}{\beta}\widehat{c}_{t+1} - r\widehat{r}_{t+1} - \omega k_{t+2} + \frac{\beta\omega}{1+\beta}k_{t+1}\} \tag{A9}$$

$$\frac{\psi^{Oh}\lambda_n}{\psi^{Oh}\lambda_n + bc^{-\eta}}(\widehat{\psi}_t^{Oh} + \widehat{\lambda}_{nt}) - \frac{\eta bc^{-\eta}}{\psi^{Oh}\lambda_n + bc^{-\eta}}\widehat{c}_t = -\zeta\widehat{l}_t \tag{A10}$$

$$\lambda_u\widehat{\lambda}_{ut} = \beta E_t\{\psi^{Ih}\lambda_n\widehat{\lambda}_{nt+1} + \psi^{Ih}[\lambda_n - \lambda_u]\widehat{\psi}_{t+1}^{Ih} + \lambda_u[(1-\mu) - \psi^{Ih}]\lambda_{ut+1} + \phi\zeta l^{-\zeta}\widehat{l}_{t+1} - b\eta c^{-\eta}\widehat{c}_{t+1}\} \tag{A11}$$

$$\lambda_n\widehat{\lambda}_{nt} = \beta E_t\{wc^{-\eta}\widehat{w}_{t+1} - \eta wc^{-\eta}\widehat{c}_{t+1} + (1-\sigma)\lambda_n\widehat{\lambda}_{nt+1} + \sigma\lambda_u\widehat{\lambda}_{ut+1} + \phi\zeta l^{-\zeta}\widehat{l}_{t+1}\} \tag{A12}$$

$$\widehat{c}_t = E_t\widehat{c}_{t+1} - \frac{1}{\eta}(\widehat{R}_t - E_t\pi_{t+1}) \tag{A13}$$

$$\widehat{y}_t = \varphi\widehat{k}_t + (1-\varphi)[\widehat{z}_t + \widehat{n}_t] \tag{A14}$$

$$\widehat{\psi}_t^{If} = \widehat{m}_{It} - \widehat{v}_t \tag{A15}$$

$$\widehat{\psi}_t^{Of} = \widehat{m}_{Ot} - \widehat{v}_t \tag{A16}$$

$$\frac{1}{\psi^{If}+\psi^{Of}}\left[\psi^{If}\widehat{\psi}_t^{If} + \psi^{Of}\widehat{\psi}_t^{Of}\right] + \eta\widehat{c}_t = \eta E_t\widehat{c}_{t+1} + \frac{\psi^{If}+\psi^{Of}}{\varkappa}\beta(1-\varphi)x\frac{y}{n}E_t[\widehat{n}_{t+1} - \widehat{x}_{t+1} - \widehat{y}_{t+1}] + \beta\frac{w(\psi^{If}+\psi^{Of})}{\varkappa}E_t\widehat{w}_{t+1} + \frac{1-\sigma}{\psi^{If}+\psi^{Of}}E_t\left[\psi^{If}\widehat{\psi}_{t+1}^{If} + \psi^{Of}\widehat{\psi}_{t+1}^{Of}\right]$$

$$w\widehat{w}_t = (1-\vartheta)(1-\varphi)x\frac{y}{n}[\widehat{x}_t + \widehat{y}_t - \widehat{n}_t] + (1-\vartheta)\frac{\varkappa}{\psi^{If}+\psi^{Of}}\left[\psi^{Ih}\widehat{\psi}_t^{Ih} + \psi^{Oh}\widehat{\psi}_t^{Oh}\right] \tag{52}$$

$$-(1-\vartheta)\frac{\varkappa(\psi^{Ih}+\psi^{Oh})}{\psi^{If}+\psi^{Of}}\left[\psi^{If}\widehat{\psi}_t^{If} + \psi^{Of}\widehat{\psi}_t^{Of}\right] \tag{53}$$

$$\pi_t = \beta E_t \pi_{t+1} + \frac{(1-\beta\chi)(1-\chi)}{\chi}\widehat{x}_t \tag{A19}$$

$$\widehat{R}_t = \xi_\pi \pi_t + \varepsilon_t^R \tag{A20}$$

$$\widehat{r}_t = \widehat{x}_t + \widehat{y}_t - \widehat{k}_t \tag{A21}$$

$$\widehat{y}_t = \frac{c}{y}\widehat{c}_t + \frac{i}{y}\widehat{i}_t + \frac{G}{y}\widehat{g}_t + \frac{\varkappa}{y}v\widehat{v}_t \tag{A22}$$

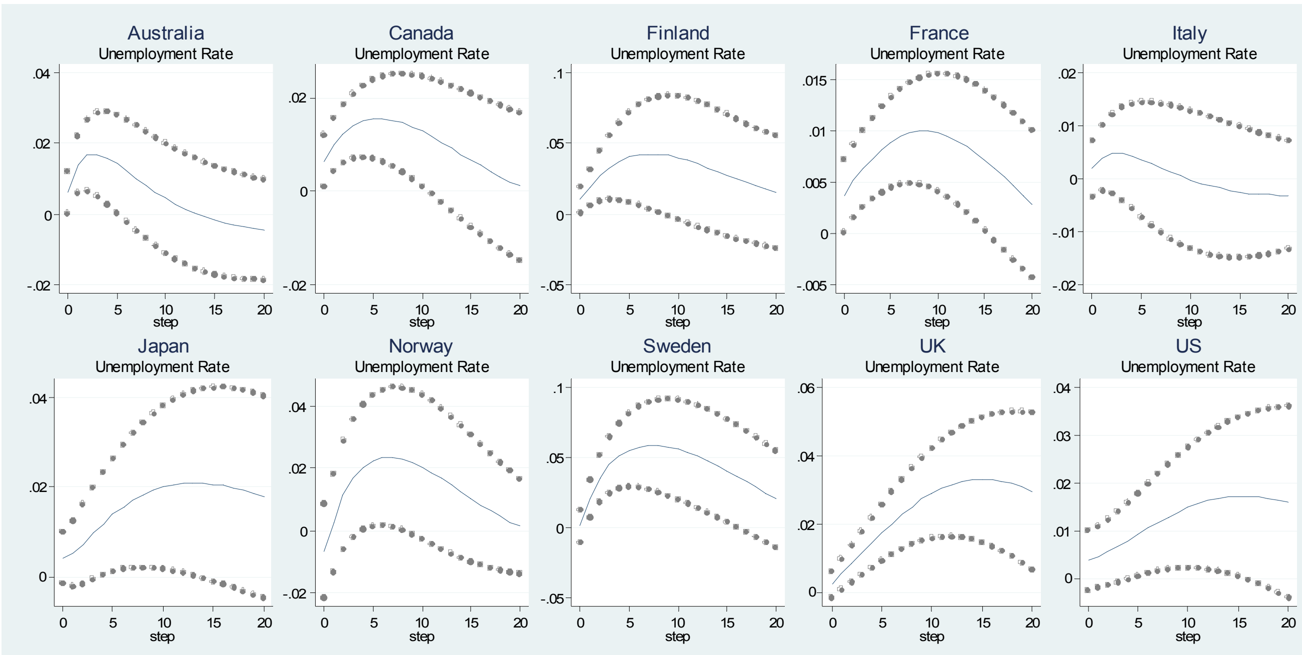
The model contains 22 equations in 22 unknowns $(n_t, m_{It}, m_{Ot}, v_t, u_{It}, u_{Ot}, \psi_t^{Ih}, \psi_t^{If}, \psi_t^{Oh}, \psi_t^{Of}, k_t, i_t, w_t, l_t, c_t, r_t, \lambda_{nt}, \lambda_{ut}, \pi_t, R_t, y_t, x_t)$ and we solve it using the generalized Schur form.

Table 1: Parameter values

β	discount factor	0.99
$1/\eta$	intertemporal elasticity of substitution	2
$1/\zeta$	elasticity of labor supply	0.25
Φ	utility of leisure parameter	2.35
$\alpha=\theta$	relative bargaining power	0.6
ρ_m^I	elasticity of new matches with respect to number of insiders	0.9
ρm^O	elasticity of new matches with respect to number of outsiders	0.7
σ	separation rate	0.09
μ	probability of becoming outsider	0.1
κ/y	cost of vacancies as a % to GDP	0.01
δ	depreciation rate	0.025
ω	capital adjustment costs	5.5
φ	capital share	0.36
ψ	probability of not changing price	0.75
$\varepsilon/(\varepsilon\text{-}1)$	gross steady state markup	1.2
ρ_g	persistence of government spending shock	0.75

Figure 1. The Effect of Government Expenditure Shocks on the Unemployment Rate

Panel A: Baseline VAR Ordering with Government Expenditures First



Panel B: Robustness VAR Ordering with Government Expenditures Last

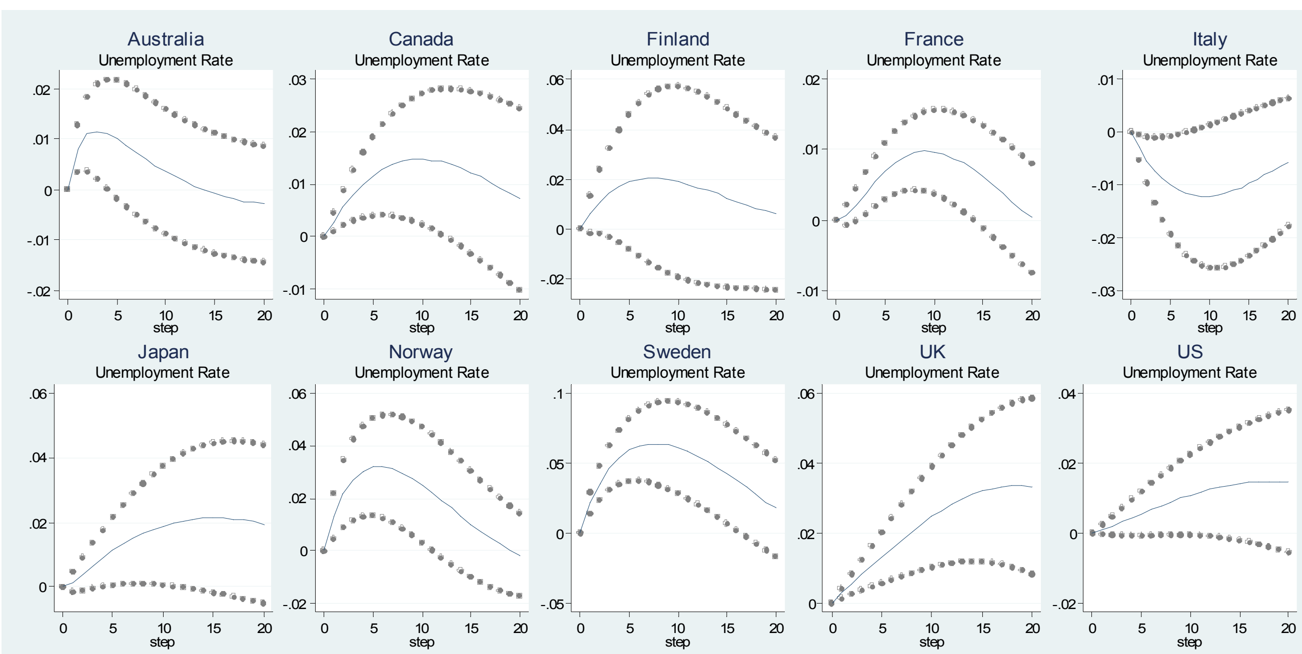
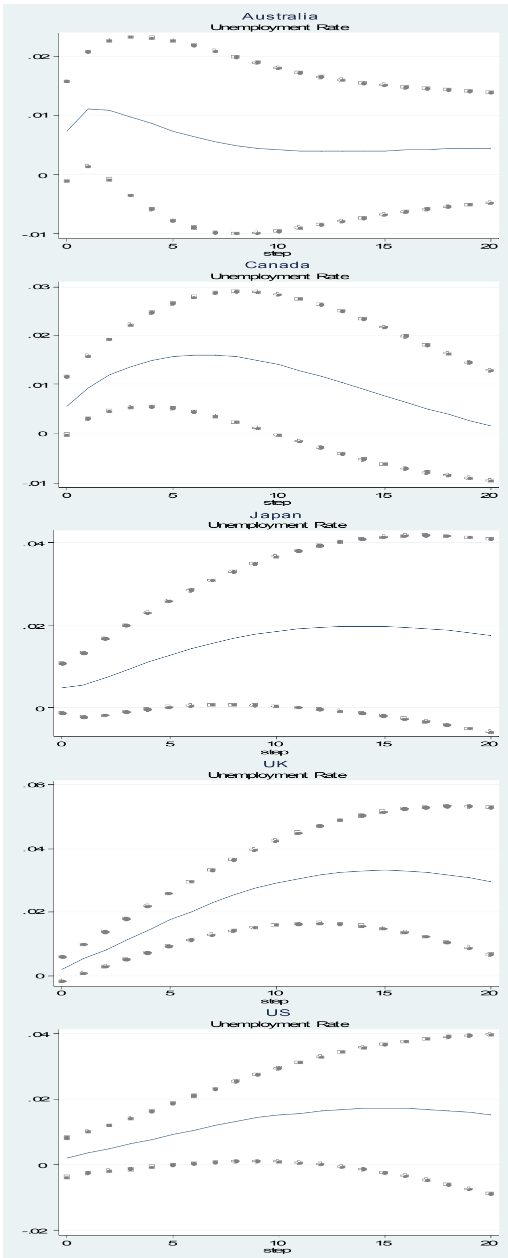


Figure 2. The Effect of Government Expenditure Shocks on the Unemployment Rate
(Controlling for Tax Revenues)

Panel A: Baseline VAR Ordering
With Government Expenditures Last



Panel B: Robustness VAR Ordering
With Government Expenditures Last

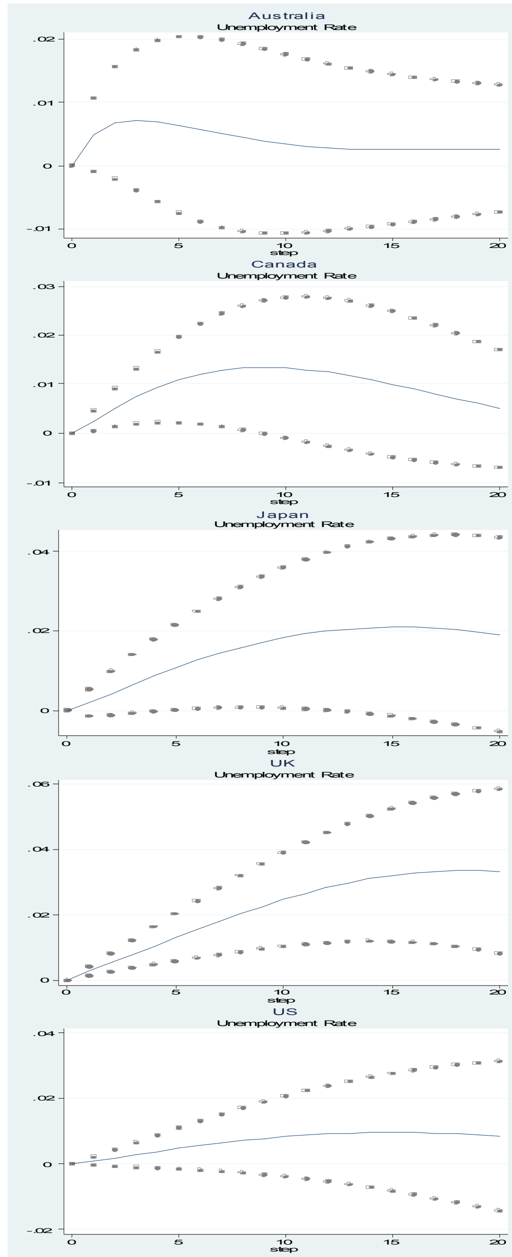


Figure 3. The Effect of Government Expenditure Shocks on the Unemployment Rate
(Ramey-Shapiro War Dummy Approach)

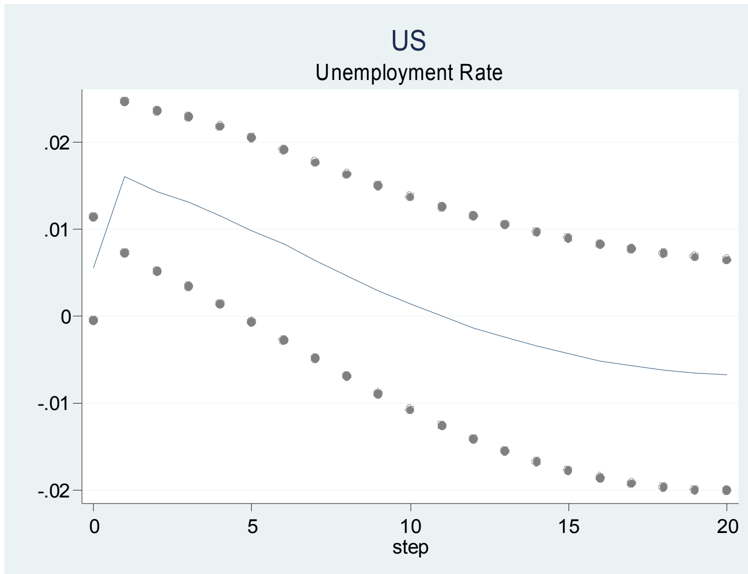


Figure 4. The Effect of Government Expenditure Shocks on the Unemployment Rate
(Different Sub-Samples Including the Pre-1980 Period)

Australia Canada US

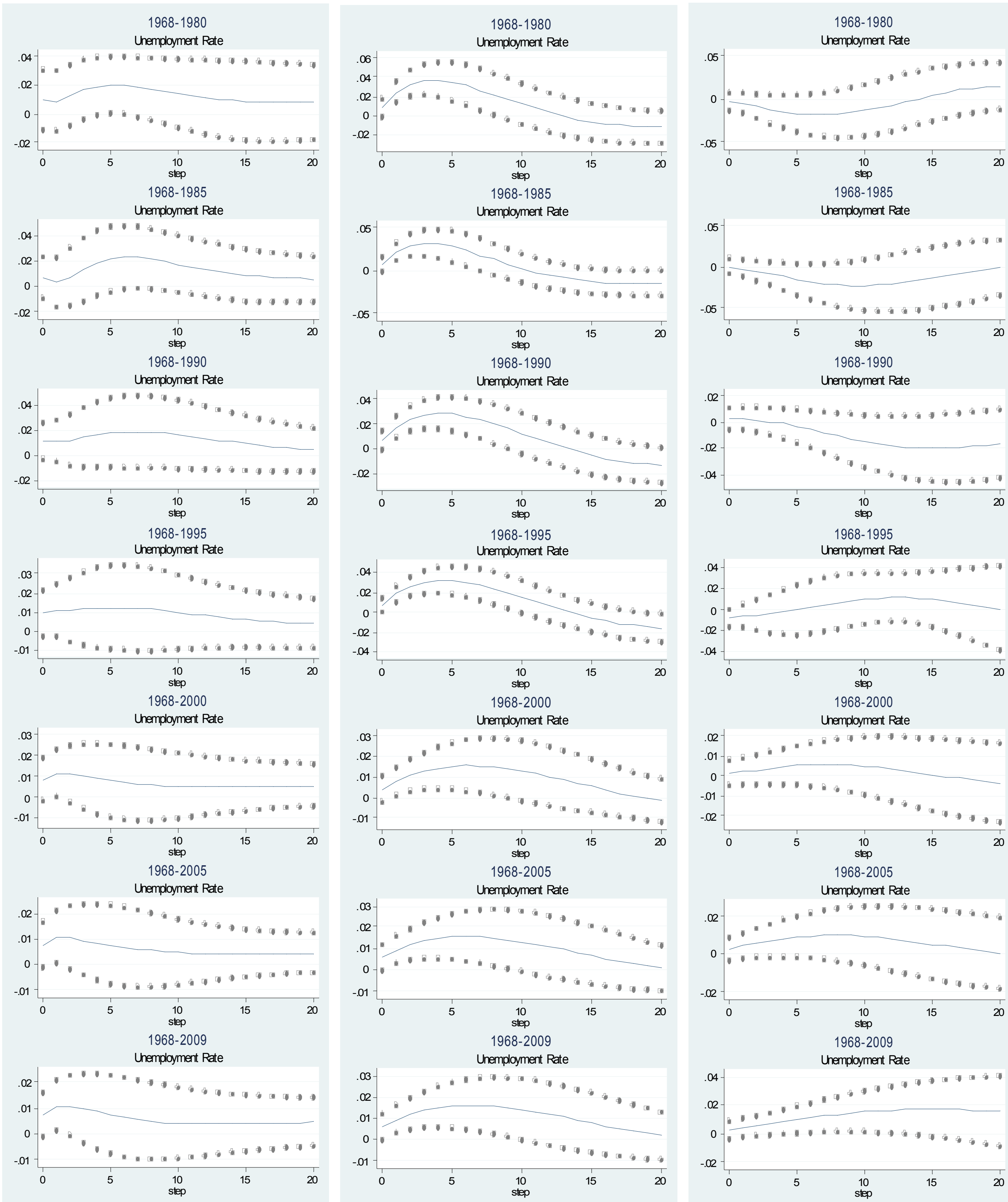
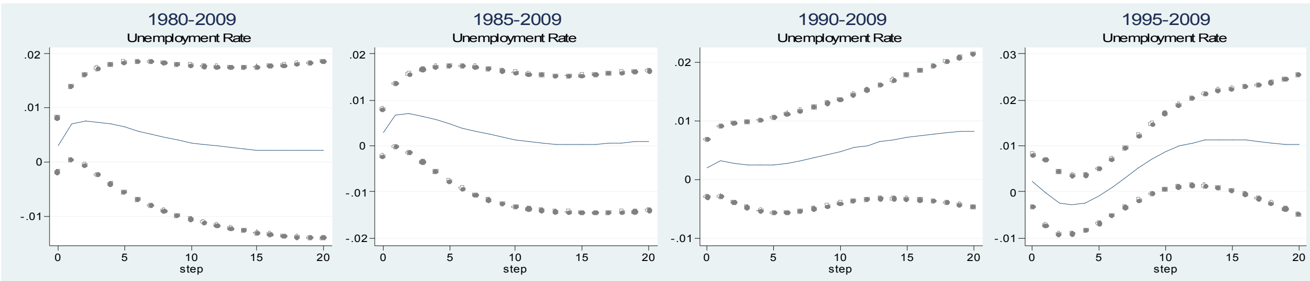


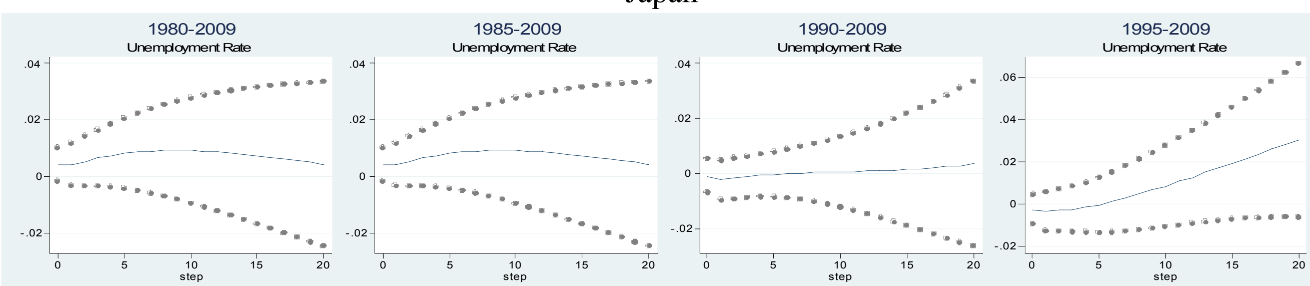
Figure 5. The Effect of Government Expenditure Shocks on the Unemployment Rate
(Different Sub-Samples Excluding the Pre-1980 Period)

Australia

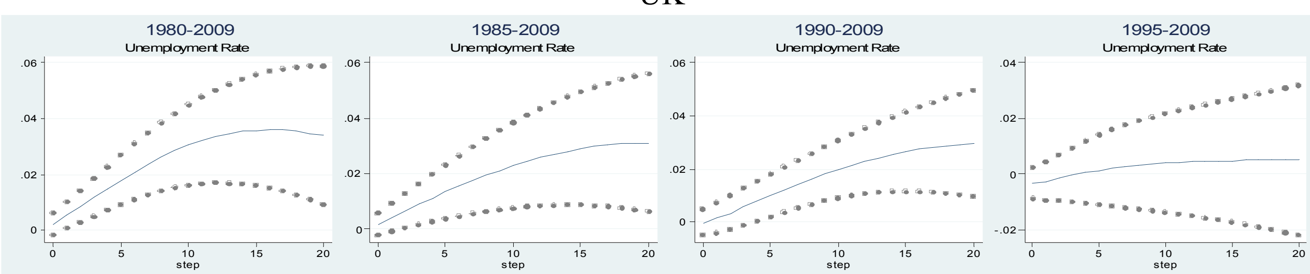


Canada

Japan



UK



US

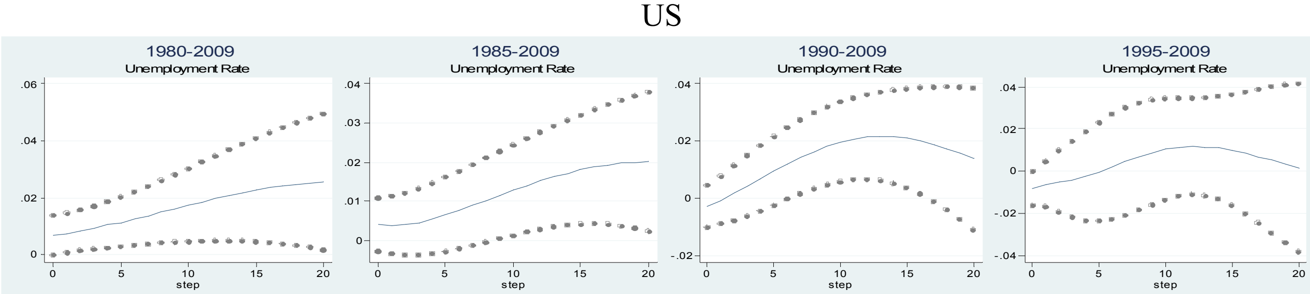


Figure 6. The Effect of Government Expenditure Shocks on Output, Consumption, and Investment

Panel A: Output | Panel B: Consumption | Panel C: Investment

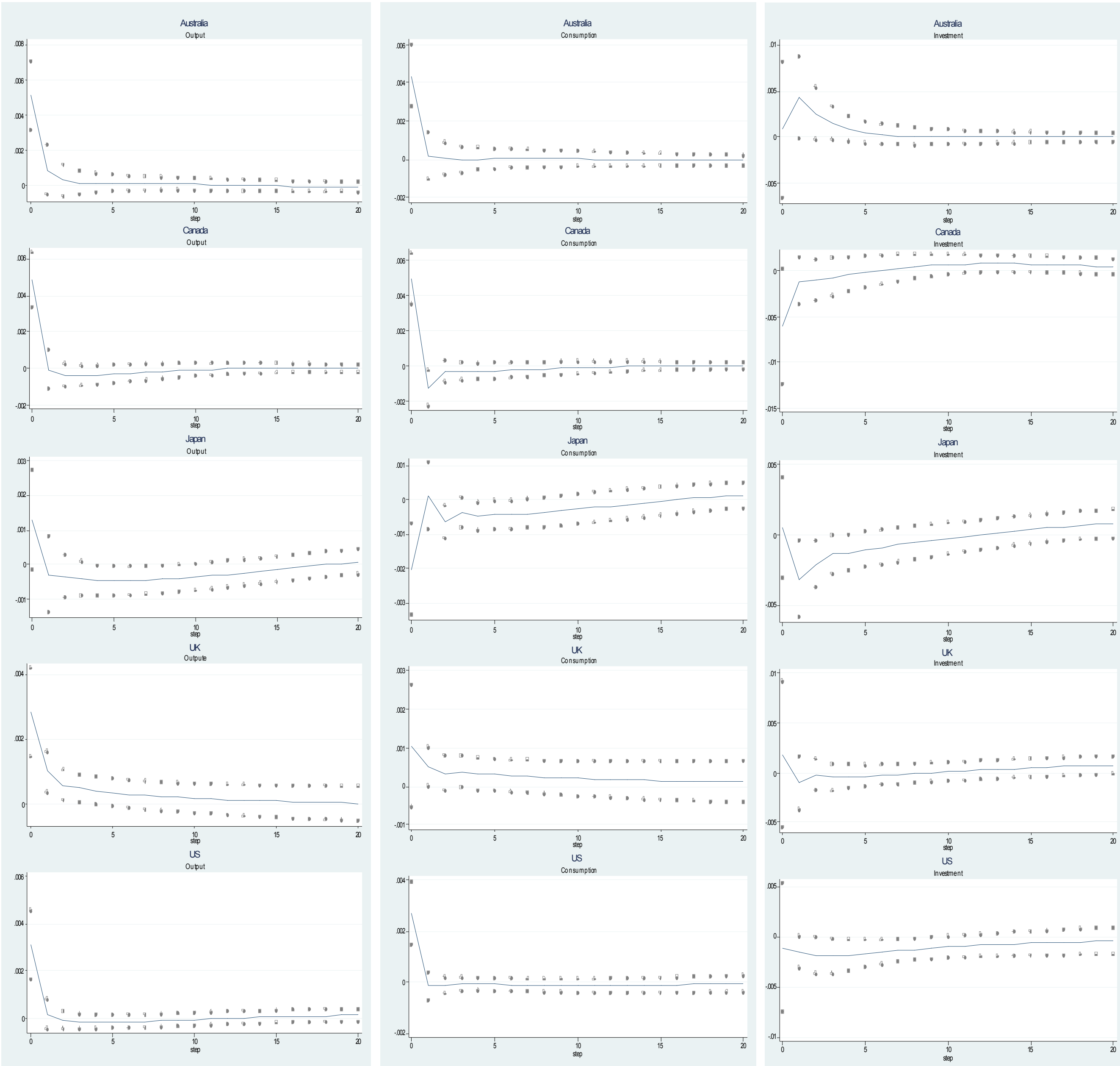


Figure 7. The Effect of Government Expenditure Shocks on Real Wages, Employment, and Labor Force Participation

Panel A: Real wage   Panel B: Employment   Panel C: Labor Participation

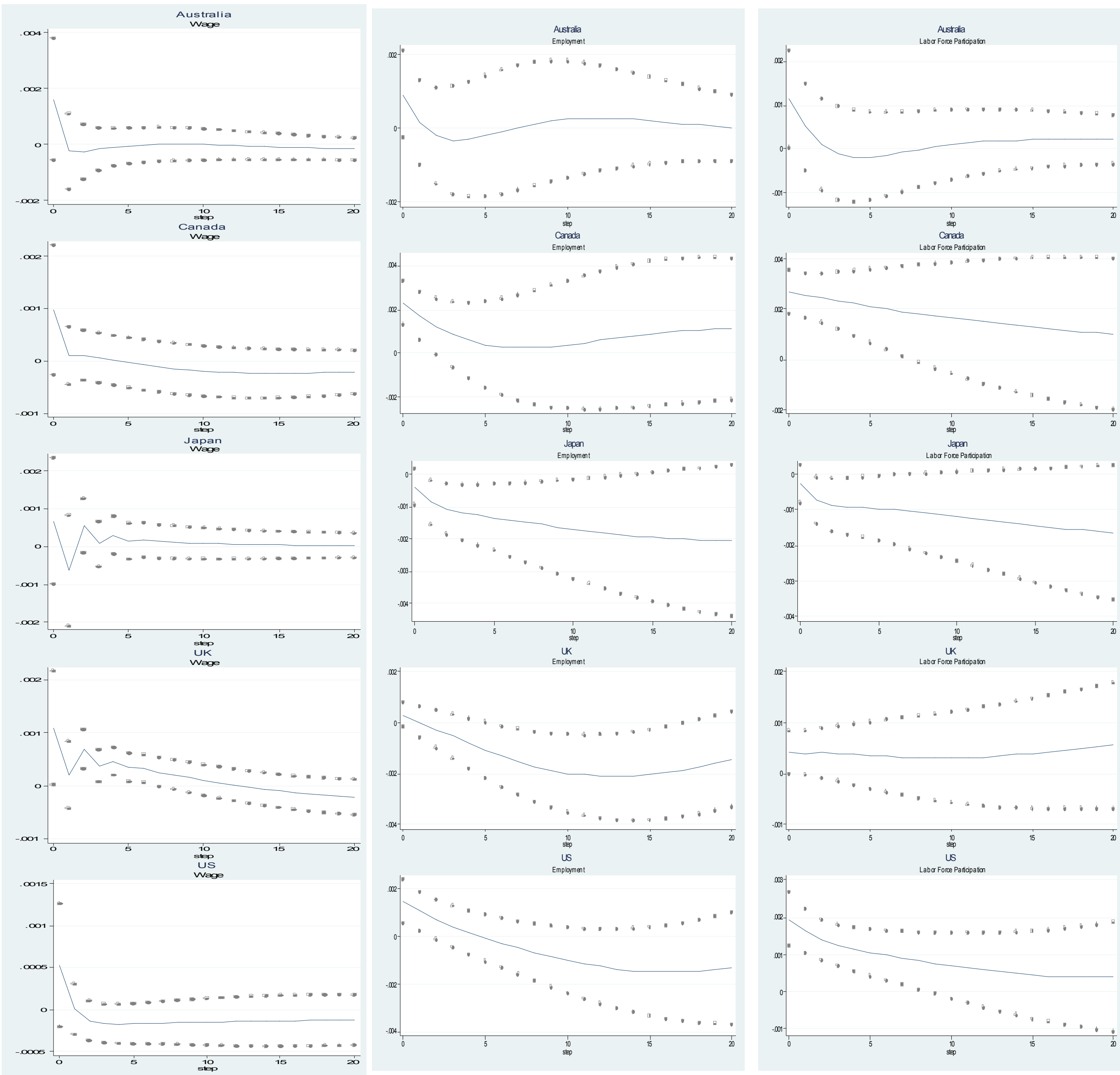
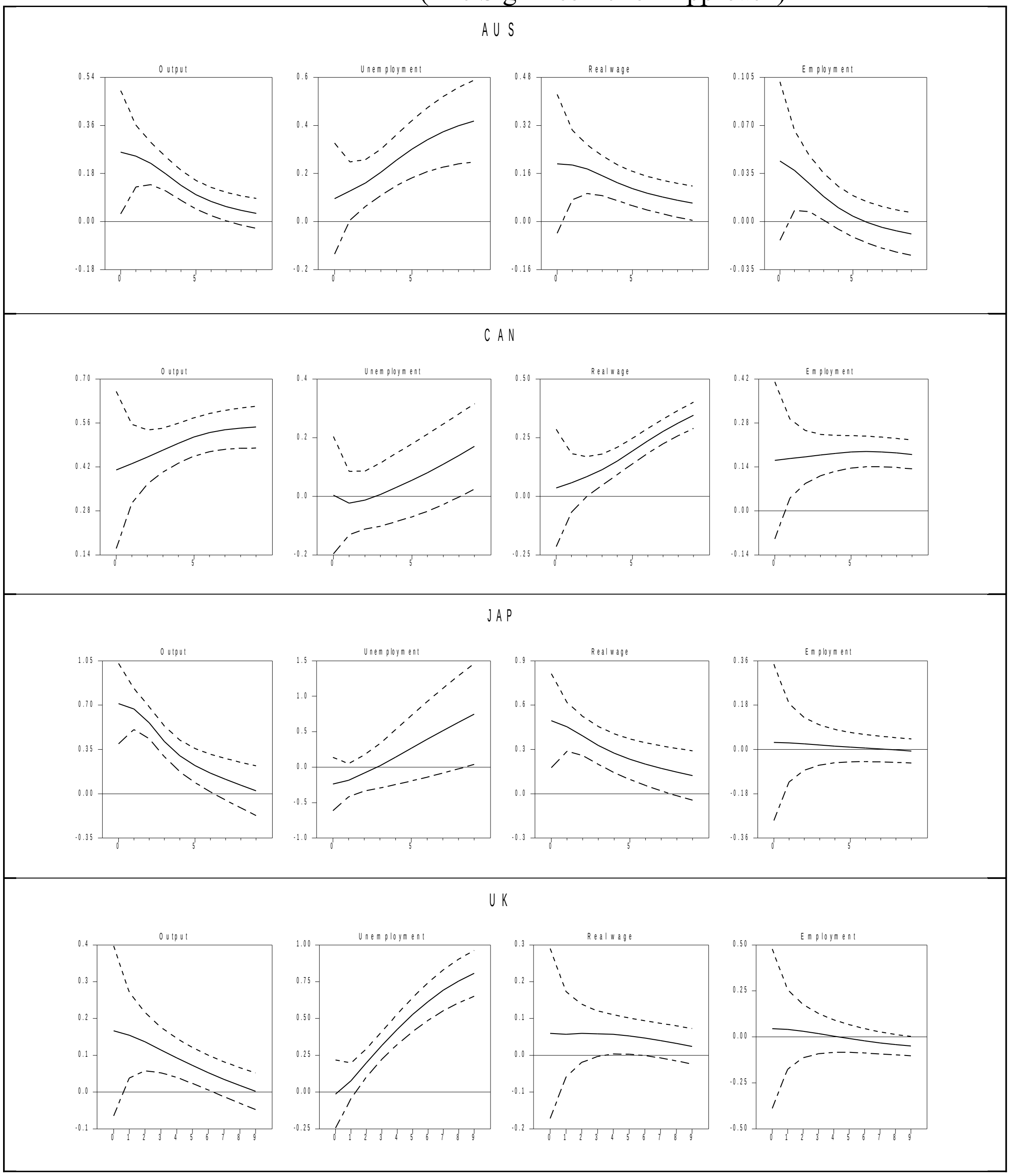


Figure 8. The Effect of Government Expenditure Shocks on the Unemployment Rate
(The Sign Restriction Approach)



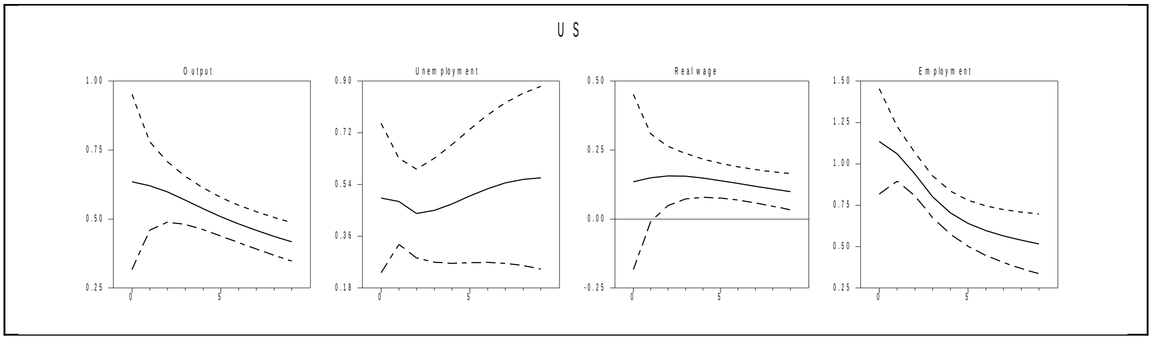


Figure 9. The Effect of Government Consumption Shocks on the Unemployment Rate
(Excluding Public Wages and Salaries)

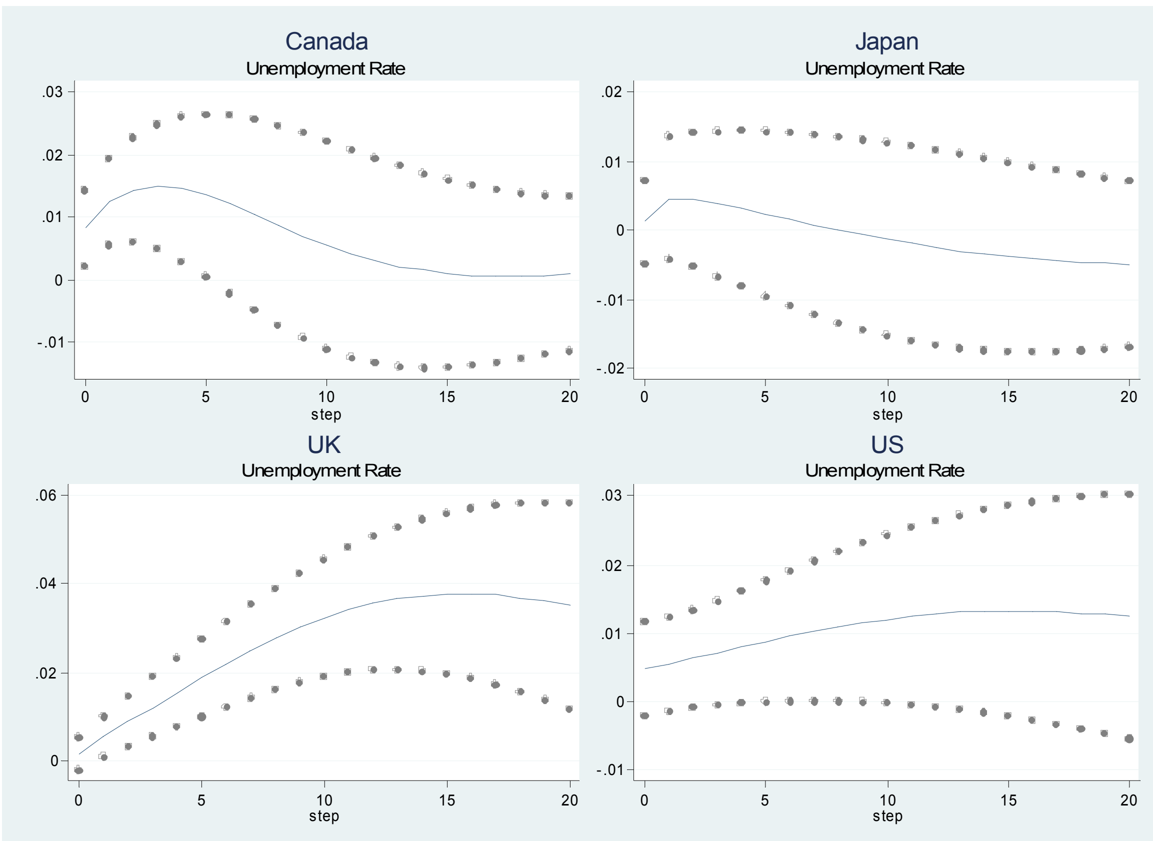


Figure 10. The Effect of Government Expenditure Shocks on the Unemployment Rate

Panel A: Controlling for Stock Prices Panel B: Controlling for Oil Price

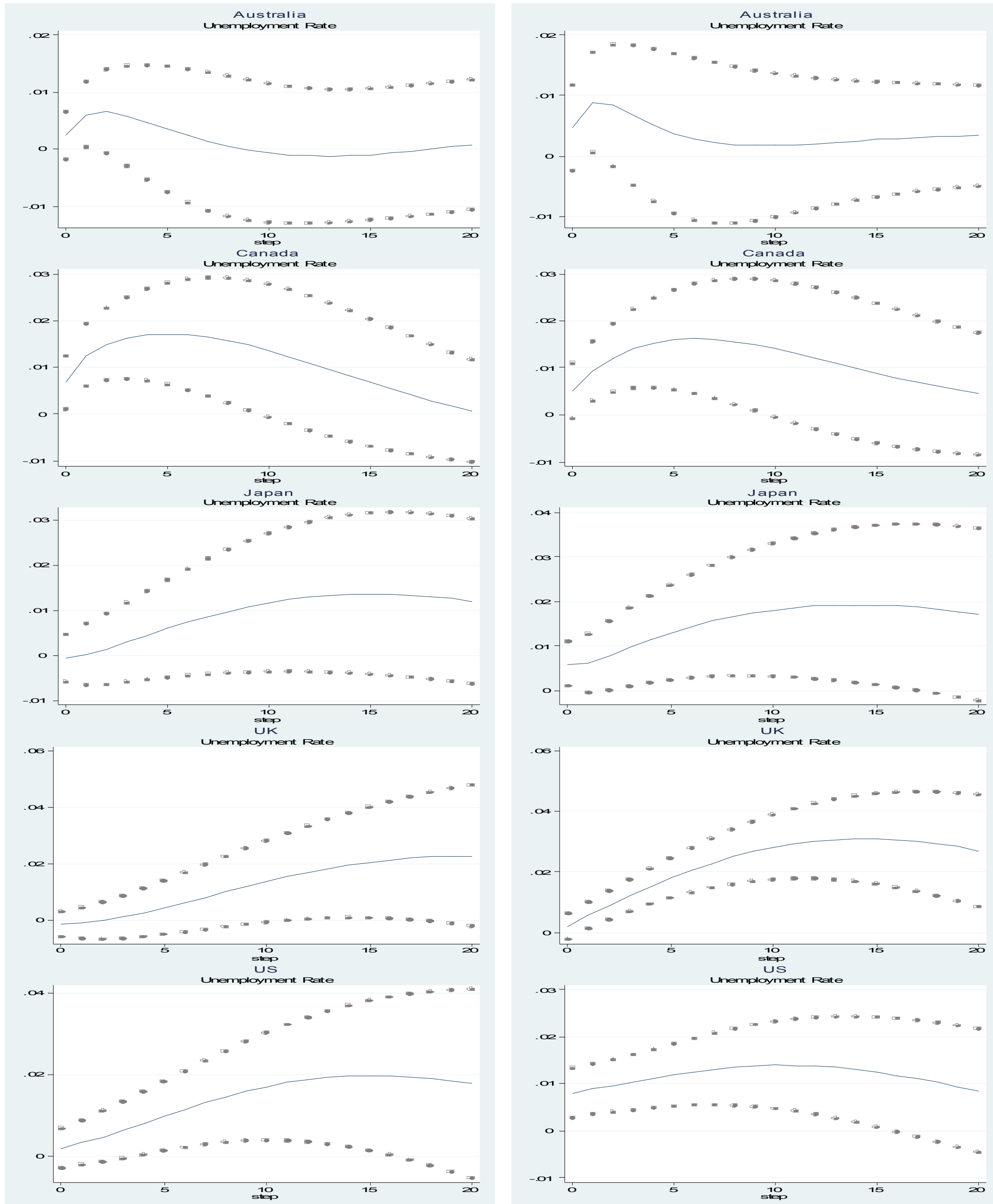


Figure 11. Theoretical Impulse Responses: Benchmark Economy

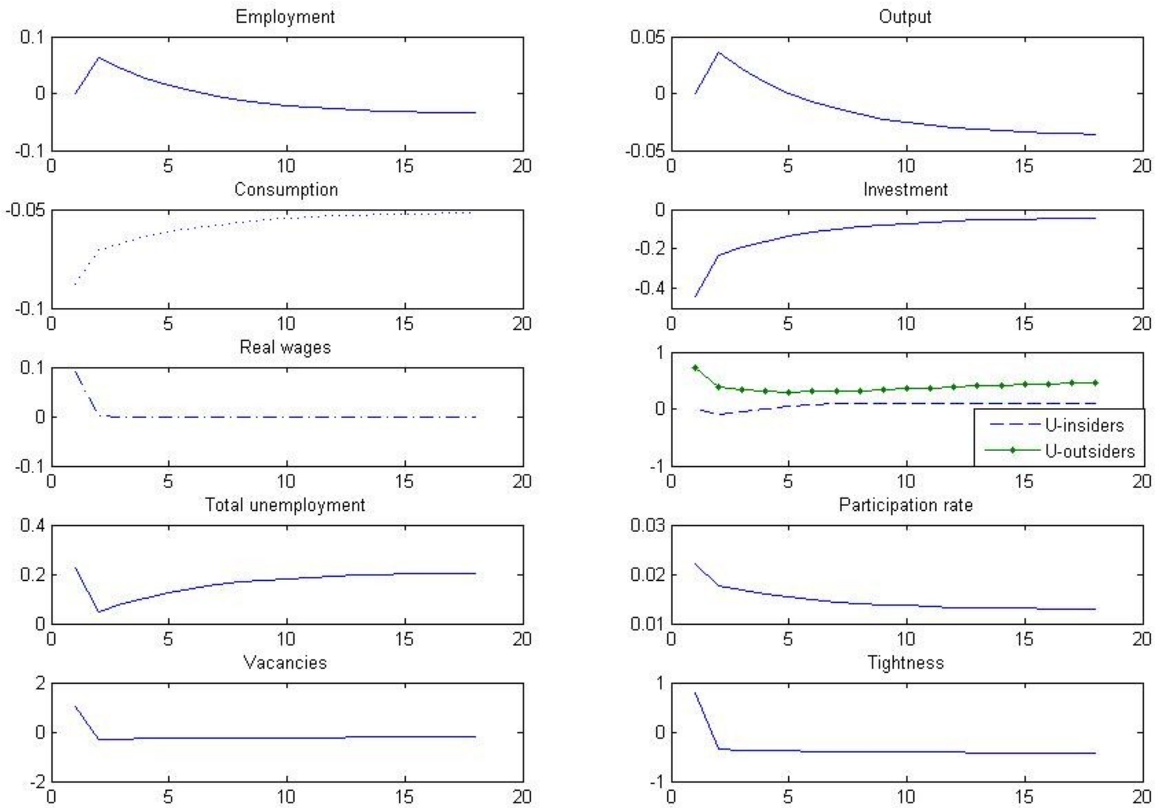


Figure 12. Theoretical Impulse Responses: Flexible vs. Sticky Prices

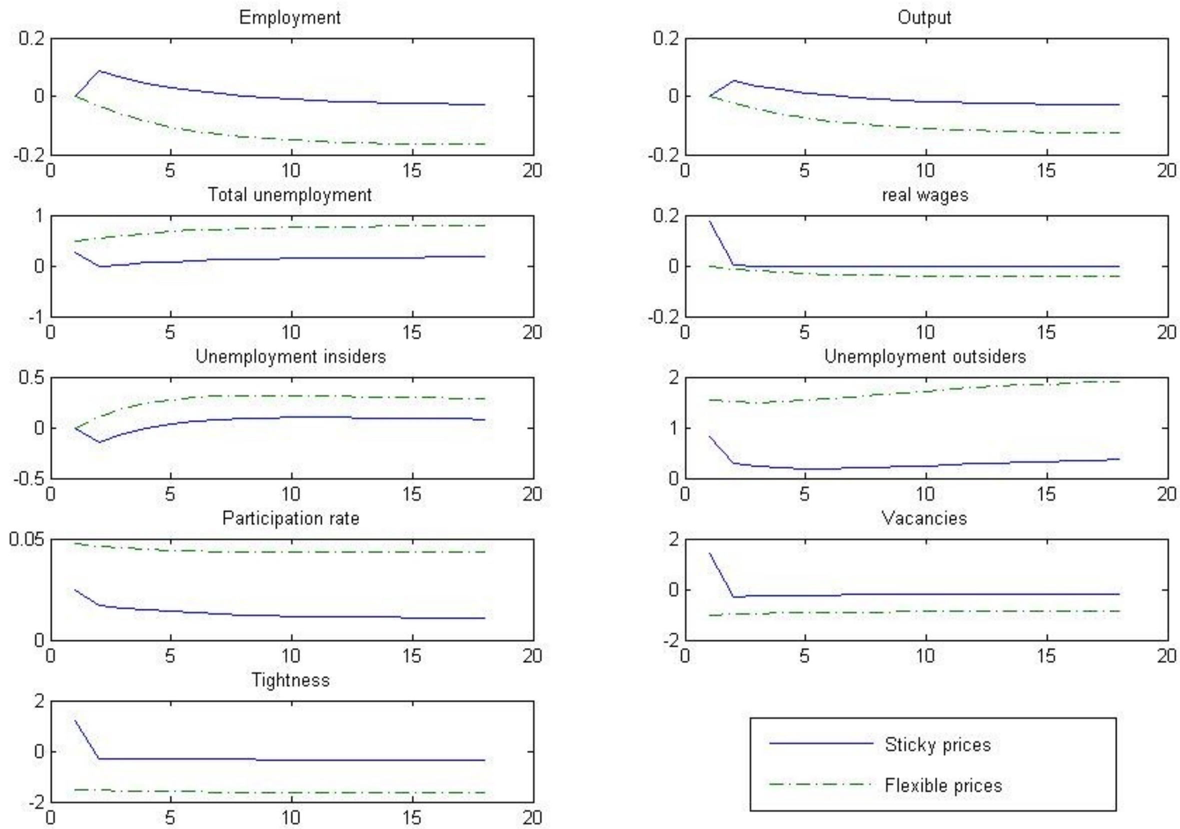


Figure 13. Theoretical Impulse Responses: Participation Margin

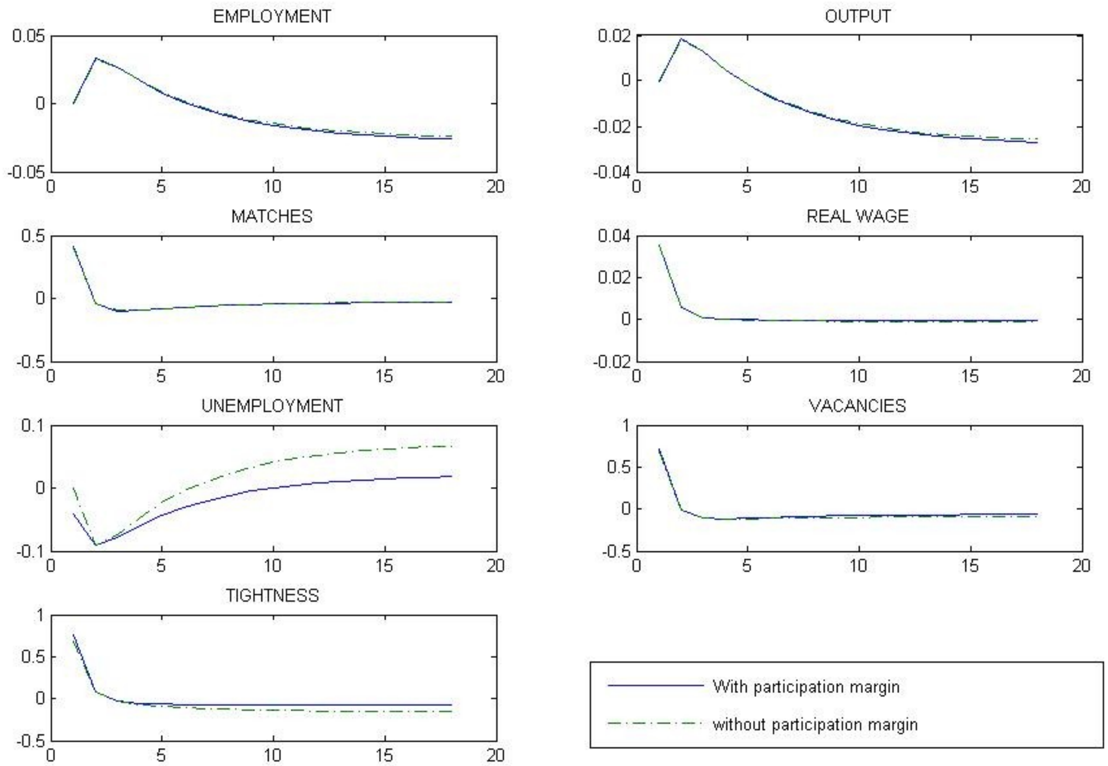


Figure 14. Participation margin and the labor supply elasticity

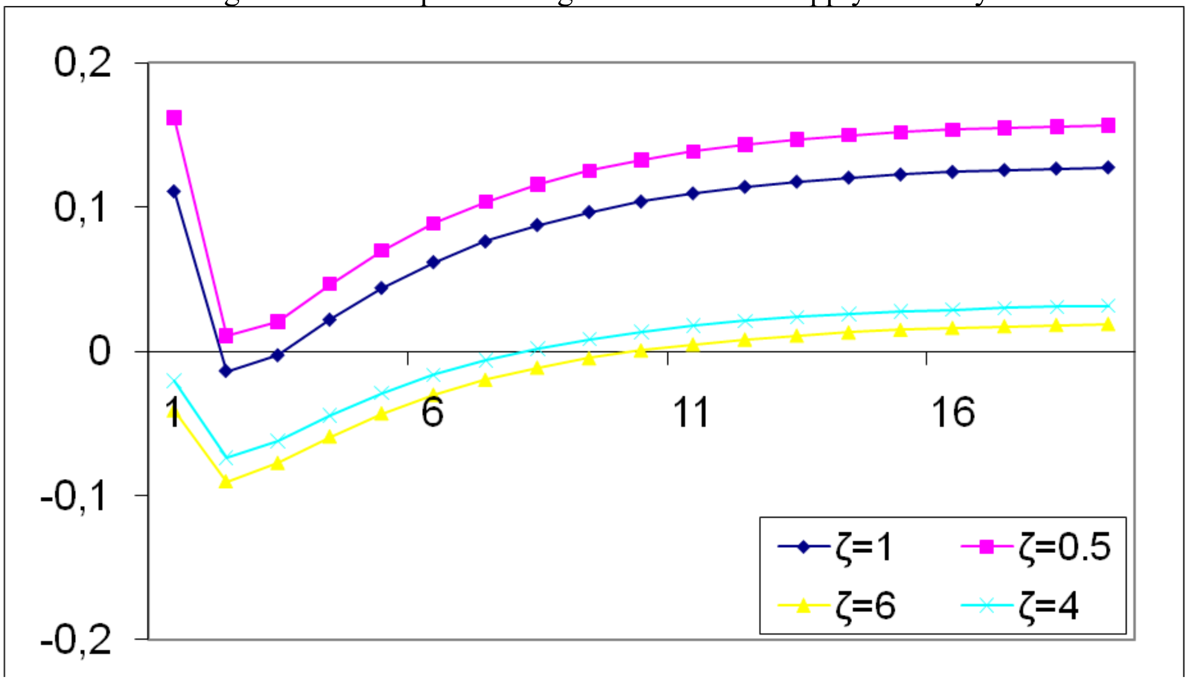


Figure 15. Sensitivity analysis

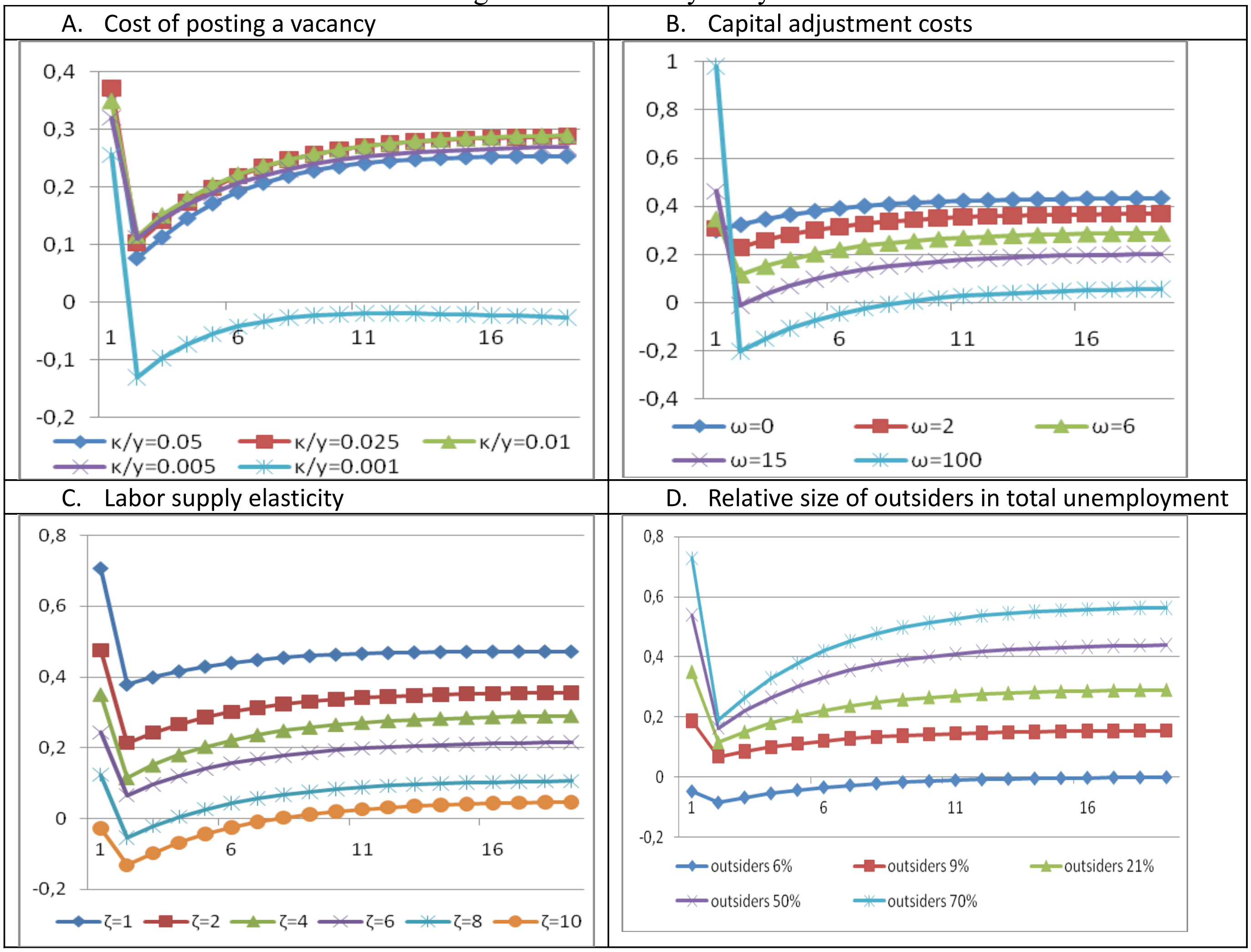
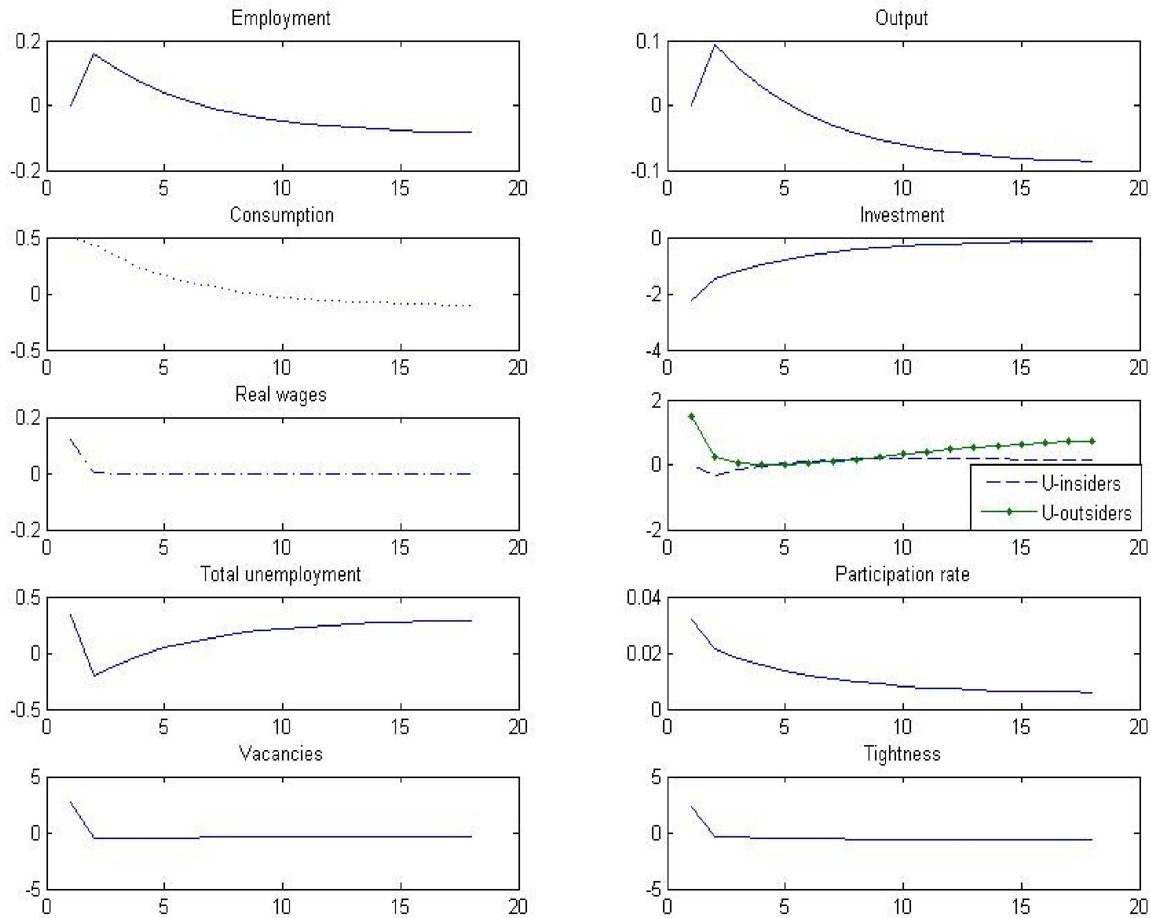


Figure 16. Complementarities of private and public consumption



Chapter 3: Economic Growth and the Rise of Political Extremism (Joint with Hans Grüner)

1 Introduction

Over the last decades, political parties with extreme platforms challenged more moderate incumbents in many western democracies (see Figures 1 and 2). Such extreme platforms rarely gained large vote shares, but frequently their success affected the political positions of more moderate parties - and so political outcomes. This paper analyses the impact of economic growth on the support for extreme political platforms. We provide a theoretical argument in favor of growth effects (as opposed to level effects) on the support for parties with extreme political platforms and we empirically investigate the relationship between growth and extreme votes.

It is not straightforward to define – in economic terms – what an extreme political platform is. Our view of extremism applies to countries in which there is some democratic competition amongst a few long term incumbent parties and where competition is limited to only a small subset of the entire policy space. In many democratic countries, there seems to exist a broad consensus about what constitutes such a set of decent policies - i.e. policies that only redistribute resources among the members of society within certain bounds.[1] In this context, we call an entrant's political platform extreme if it includes major differences in the distribution of resources compared to standard policies. In practice, such extreme political platforms often propose to redistribute resources away from specific subgroups of society (such as the rich, ethnic minorities, or citizens of specific regions).

Our analysis is based on the observation that extreme parties are frequently perceived to create more uncertainty about future policy outcomes than established parties. One reason for this is that extreme parties often have little or no government experience. Another reason may be that, once a political movement based on an extreme platform has come to power, the political elite may define new - and different subgroups of society that become the subject of redistribution. Historically, many regimes that were based on an extreme political agenda had the feature that some groups of society - be it ethnical, educational or professional - were stigmatized and suffered from

[1] For related theoretical analyses see Artale and Grüner (2003), and Grüner (2007).

redistribution or oppression.[2]

The choice of such a regime comes along with a cost when no group that benefits today can be really sure that this will stay so in the future. In the long run, this creates an income risk for all citizens and a trade-off between short-run gains from redistribution and long-run losses due to increases in income uncertainty. Economic growth increases the cost of uncertainty and so increases support for a moderate regime.

In the first part of the paper we develop a simple game theoretic model that further analyzes these effects. The purpose of the model is twofold. First, it shall give reasons for why economic growth and not just the level of income may have an impact on the support for moderate political regimes. Second, it shall provide testable comparative static results about the determinants of political radicalism.

In our model, extreme political parties offer short-run gains from redistribution to many individuals. However, the same individuals also face long-run losses due to more instability and higher income risk. Only sufficiently poor agents are in favor of extreme policy platforms. The model permits a comparative static analysis with respect to several variables of interest. The growth rate is associated with a higher cost of future income risk. This reduces the number of voters in favor of extreme policies. Similarly, a higher discount factor raises the vote share of moderate platforms. The share of stigmatized agents has ambiguous effects on the support for the moderate regime. On the one hand, it increases revenues from redistribution, on the other hand, stigmatized agents favor moderate policies. Moreover, the scope for future expropriation may also be affected. Economic inequality raises the support for redistribution and it also affects the effect of changes in economic growth.

An important prediction of our model is that the effects of economic growth on the support for an extreme political party depends on the perceived likelihood that this party will generate unstable policies that affect different ethnic, regional or religious subgroups of society over time. If policies are perceived as stable - in the sense that the same groups of society remain priviledged, political support of this party is unaffected by growth.

In the empirical part of our paper we construct a panel dataset for 16 OECD countries that includes survey-based measures of political support for

[2] Frequently, extreme political parties with a small membership basis attract a large number of dissatisfied voters. The interaction of these voters and the party members is hard to predict. This adds to the uncertainty about the political consequences of an electoral outcome.

right-wing/nationalist parties and communist parties. We use this data to approximate the support for extreme political platforms. We apply rigorous panel data techniques to estimate the impact that economic growth has on the share of voters who favor such platforms. Any attempt to investigate the relationship between growth and the support for certain policies is plagued by causality problems - support for different policies is likely to shape institutions and institutions are likely to affect growth. We address this causality issue by using instrumental variable techniques and panel fixed effects regressions. Specifically, we use both system-GMM estimation as well as international oil price shocks as instrumental variables to deal with endogeneity issues. We deal with unobservable cross-country heterogeneity and common year shocks by using country and time fixed effects.

Our main finding is a negative and significant effect of real per capita GDP growth on the support for extreme political parties. At the same time, our analysis also makes clear that even major changes in the GDP per capita growth rate will most likely not change the political outcome in any of the OECD economies substantially. According to our estimates, a one percentage point drop in real per capita GDP growth would on average increase the share of extreme right-wing political parties by roughly one percentage point. In most economies this is unlikely to have any lasting impact on the political outcomes.

It is particularly noteworthy that there is a differential effect of growth on left-wing and right-wing extremism. There is a clear effect on the support for extreme right-wing parties whereas we find little evidence on the support for communist parties. To the extent that communist parties mainly wish to redistribute from the rich to the poor, this is in line with our theoretical predictions.

Our paper is related to a literature that investigates the relationship of economic development and political outcomes. For a long time, social scientists have argued that income and democracy go hand in hand. Two different kinds of theoretical arguments have been made in favor of a positive relationship between income and democracy. The first class of explanations concerns a possible causality that goes from income to democracy.[3] The most popular one is that a higher income level enables an emerging middle class to success-

[3]See for example Geddes (1999), Przeworski et al. (2000), Glaeser et al. (2004), Acemoglu et al. (2008, 2009), Brückner and Ciccone (2008), or Papaioannou and Siourounis (2008) among others. For earlier contributions, see Lipset (1959) or Huntington (1991).

fully fight for political emancipation.[4] The second set of arguments concerns the inverse causal direction. According to this view, democracy has a positive impact on economic freedom and so creates a higher living standard. A synthesis of both views has recently been proposed by Persson and Tabellini (2009). They argue, that voters learn from the economic performance of their political system. Citizens are only willing to defend democracy if they believe in its economic benefits. A switch from democracy to an autocratic regime is more likely when the system performs poorly in economic terms. This implies that old democracies are likely to have higher levels of GDP whereas new democracies can start with a low level of GDP.

Most of the above arguments focus on the relationship between the level of output and democratic institutions. However, some economists also argue that economic growth is another independent and major determinant of the support for and development of a democratic political system. This point has recently been raised by Benjamin Friedman (2003). Friedman argues that only a continuous improvement of individual living standards provides the ground for a sound functioning of a democratic system and for the development of a more open political system. One of the reasons Friedman gives why individuals are more content with the political system if they experience improvements of their living standards is that individual well being is linked to income growth and not just the level of income.

The remainder of the paper is organized as follows. Section 2 introduces the theoretical model. Sections 3 and 4 describe the dataset and estimation strategy. Section 5 presents the main empirical results. Section 6 concludes.

2 A simple theoretical framework

2.1 The moderate regime

Consider a population of $i = 1, ..., n$ individuals who live in periods $t = 0, \ldots, \infty$. In every period, the economy is in one of two possible political

[4] There is little theoretical or empirical work on the relationship of growth and voting outcomes. One exception is De Neve (2010), who attempts to relate the US median voter's preference for the size of the government sector to economic growth. In his model agents only derive utility from changes in private and public consumption. With an appropriate utility function, all voters prefer a higher tax rate when income growth is higher. The model has nothing to say though about the support for extreme political positions.

regimes, the moderate (M) and the extreme one (E). In regime M, all individuals have a given income, $\tilde{y}_{it}$, that grows with a constant growth rate:

$$\begin{aligned} \tilde{y}_{it} &= g^t \tilde{y}_{i0}, \text{ with } g > 1, \text{ and} && (1) \\ \frac{1}{n}\sum_{t=1}^{n} \tilde{y}_{i0} &= \bar{y}_0. && (2) \end{aligned}$$

An individual's income under the moderate regime should be thought of as the market income corrected through "standard" redistributive measures such as a progressive income tax system.

All individuals are risk averse and care about discounted utility derived from net income y_t. They maximize the expected value of

$$\sum_{t=0}^{\infty} \delta^t \mathbf{u}\left(y_{it}\right), \tag{3}$$

with $u'\left(y_t\right) > 0, u''\left(y_t\right) < 0$. More specifically, in order to obtain closed form solutions, we assume that

$$\mathbf{u}\left(y_{it}\right) = \mathbf{y}_{it}^{\alpha}. \tag{4}$$

Discounted expected utility is given by

$$\mathbf{U}^M := \sum_{t=0}^{\infty} \delta^t \mathbf{u}\left(y_{it}\right) = \sum_{t=0}^{\infty} \delta^t \left(g^t \tilde{y}_{i0}\right)^{\alpha} = \sum_{t=0}^{\infty} \left(\delta g^{\alpha}\right)^t \tilde{\mathbf{y}}_{i0}^{\alpha} = \frac{1}{1-\delta g^{\alpha}} \tilde{\mathbf{y}}_{i0}^{\alpha}. \tag{5}$$

In regime M, in each period individuals may support one of the two regimes in a vote. Either they support the existing regime M or they vote for regime E. In what follows we consider an extreme case where this policy turns the system into a persistent political regime that is characterized by high income uncertainty.

2.2 Regime E

At the beginning of each period, nature randomly selects a subset S of the $s \cdot n$ individuals that are stigmatized. In each period, every individual knows, whether he or she belongs to the set S or not. In regime E all incomes $\tilde{y}_i$ are collected by the state (who also observes S) and redistributed evenly across all individuals who are not stigmatized. Therefore net incomes in period t are

$$\mathbf{y}_{it}=\begin{cases} \frac{1}{1-s}g^t\bar{y}_0 & \text{if } i\notin S \\ 0 & \text{if } i\in S \end{cases}. \tag{6}$$

For simplicity, we assume that agents have no choice in an extremist regime; i.e. such a regime persists. Permitting the return to the moderate regime would not affect our results. Discounted expected utility of agents in $N\setminus S$ in an regime E, beginning at t=0, is:

$$\begin{aligned} U^E := \quad & u\left(\frac{\bar{y}_0}{1-s}\right)+\sum_{t=1}^{\infty}\delta^t(1-s)\,u\left(\frac{g^t\bar{y}_0}{1-s}\right) && (7)\\ = \quad & su\left(\frac{\bar{y}_0}{1-s}\right)+\sum_{t=0}^{\infty}\delta^t(1-s)\,u\left(\frac{g^t\bar{y}_0}{1-s}\right) && (8)\\ = \quad & s\left(\frac{\bar{y}_0}{1-s}\right)^{\alpha}+(1-s)^{1-\alpha}\frac{\bar{y}_0^{\alpha}}{1-\delta g^{\alpha}}. && (9) \end{aligned}$$

2.3 Equilibrium

A strategy of an agent maps the history of the game into a voting decision. Without restricting generality, we consider the optimization problem of an agent in period 0. An agent who is not stigmatized in period 0 prefers the continuation of the status quo to an extreme political regime if

$$\begin{aligned} U^M \quad & > \quad U^E \Leftrightarrow && (10)\\ \frac{1}{1-\delta g^{\alpha}}\tilde{y}_{i0}^{\alpha} \quad & > \quad s\left(\frac{\bar{y}_0}{1-s}\right)^{\alpha}+(1-s)^{1-\alpha}\frac{\bar{y}_0^{\alpha}}{1-\delta g^{\alpha}}\Leftrightarrow && (11)\\ \tilde{y}_{i0} \quad & > \quad Y := \left((1-\delta g^{\alpha})\,s\left(\frac{1}{1-s}\right)^{\alpha}+(1-s)^{1-\alpha}\right)^{\frac{1}{\alpha}}\bar{y}_0. && (12) \end{aligned}$$

The same condition applies to all further periods. Therefore, players have the following weakly dominant strategies. All agents with initial income $\tilde{y}_{i0}\geq(<)Y$ support (oppose) regime M in all periods, independently of whether they are stigmatized in period t or not. Stigmatized agents with initial income $\tilde{y}_{i0}$

support regime M if

$$\frac{1}{1-\delta g^{\alpha}}\tilde{y}_{i0}^{\alpha} \geq 0 + \delta g^{\alpha}(1-s)^{1-\alpha}\frac{\bar{y}_0^{\alpha}}{1-\delta g^{\alpha}} \Leftrightarrow \tilde{y}_{i0} > Y' := \delta^{\frac{1}{\alpha}} g (1-s)^{\frac{1-\alpha}{\alpha}}\bar{y}_0. \tag{13}$$

Otherwise, they support regime E. Note that, for appropriate parameters δ, g, α, and s the threshold level Y is below the initial average income $\bar{y}_0$. Therefore, societies in which the median of the income distribution is below the mean need not necessarily turn into an extreme political regime. Moreover, as one can easily verify, the threshold income Y' above which stigmatized agents prefer the status quo always lies below Y if $\delta g^{\alpha} < 1$. This condition must hold for the discounted sum of utilities to exist.

2.4 Results

Our simple theoretical model produces a number of useful results.[5]

1. A higher discount factor increases support for the moderate regime because agents care more about the future income risk.

2. A higher growth rate increases support for the moderate regime because it increases the variance of future income in an extreme political regime.

3. A higher individual income raises an individual's support for the moderate regime.

4. Consider an alternative distribution of income at date zero that preserves the income ratio $\tilde{y}_{i0}/\bar{y}_0$ for all individuals. It follows from (12) and (13) that all individuals favour the moderate regime if and only if they did so under the old income distribution. Hence, ceteris paribus, the initial average income $\bar{y}_0$ does not affect the political outcome.

5. Inequality (measured by the share of individuals who earn less than Y) reduces support for the moderate regime.

6. Consider a uniform distribution of initial income with a given mean. Inequality reduces the marginal effect of growth on the support for regime M.

[5]The results follow directly from conditions (12) and (13).

7. The share of stigmatized agents in the population s has an ambiguous effect on the support for the moderate regime. If $\delta = 1/g^{\alpha}$ then a higher share s reduces the threshold for income above which agents who are not stigmatized support the moderate regime.

8. When s=0, there is no effect of growth on the support for regime E. This means that the support for a regime that merely redistributes from the rich to the poor does not change when the growth rate increases.

In our empirical analysis that follows, we mainly concentrate on the effect of economic growth on the support for extreme political platforms (the second theoretical result). We also present some first empirical evidence on the role of level effects (result 4) and the role of inequality for the marginal effect that economic growth has on the support for extreme political parties (result 6). Moreover, in relation to result 8, we compare the effects of economic growth on the support for left-wing and right-wing parties.[6]

3 Description of the OECD Vote Share Dataset

We constructed a semi-annual panel dataset comprising 16 OECD countries for the period 1970-2002.[7] Our main measure for the rise of extreme political parties is from Eurobarometer.[8] Eurobarometer conducted from 1970

[6] We have also made an attempt to test result 5 by looking at the cross-country correlation between measures of income inequality (as well as measures of poverty) and the support for extreme political platforms. We did not find a significant relationship, which may be due to the insufficient number of cross-country observations (16) in our OECD dataset. We have also made an attempt to test hypothesis 5 with panel data, using the labor income share as a proxy for income inequality. Our main finding was that increases in the labor income share are associated with a significant within-country decrease in the support for extreme political platforms, which is consistent with result 5. Results are not reported here for space purposes and are available from the authors upon request. Note that due to lack of data on country-specific discount and stigmatization factors, we are unable to test the other results from the model.

[7] The countries (time-period) covered in our dataset are: Austria (1994-2002), Belgium (1970-2002), Denmark (1973-2002), Finland (1993-2002), France (1970-2002), West-Germany (1970-2002), Great Britain (1973-2002), Greece (1980-2002), Ireland (1973-2002), Italy (1970-2002), Luxembourg (1973-2002), Netherlands (1970-2002), Norway (1990-1995), Portugal (1985-2002), Spain (1985-2002), and Sweden (1994-2002).

[8] The data is publicly available at http://zacat.gesis.org/webview/index.jsp.

to 2002 semi-annual surveys of individuals' voting intentions in OECD countries.[9] The question asked in the Eurobarometer survey was the following: "If there were general elections tomorrow, which party would you vote for". We then constructed three variables that proxy the support for extreme political platforms in a country-period. The first variable proxies the support for right-wing/nationalist parties. This variable is constructed by summing over all the votes given to right-wing/nationalist parties (right-wing/nationalist parties are identified according to the ZEUS party code) and dividing these votes by the total number of votes in the survey. The second variable proxies the support for communist parties. This variable is constructed by summing over all the votes given to communist parties (again identified according to the ZEUS party code) and dividing these votes by the total number of votes in the survey. The third variable proxies the total support for extreme political parties and is constructed by adding the vote shares obtained by right-wing/nationalist parties with the vote shares obtained by communist parties.

Basic summary statistics of the vote share of extreme political parties in our sample are as follows. The mean vote share of right-wing/nationalist parties is 0.016. The between-country standard deviation is 0.031 and the within-country standard deviation is 0.016. The interquantile range is [0, 0.026]. 5% of all the right-wing/nationalist vote shares are larger than 0.08 and the sample maximum is 0.15. For communist parties, the mean vote share is 0.041. The between-country standard deviation is 0.044 and the within-country standard deviation is 0.025. The interquantile range is [0, 0.071]. 5% of all the communist vote shares are larger than 0.156 and the sample maximum is 0.222.

Note that the vote share of extreme political parties is heavily positively skewed. Once we demean the vote share from its country-average and the common time fixed effect the skewness disappears however. This is shown in the kernel density plot of Figures 3 and 4.

To present also some specific examples of the empirical evolution of the vote share of extreme political parties we plot in Figures 1 and 2 time-series graphs of the right-wing/nationalist vote share and the communist vote share for 4 of our 16 OECD countries (Denmark, Italy, West-Germany, and France).

[9] The average survey size was 1088, with an interquantile range of [1000, 1049]. Note that because the surveys were taken randomly across individuals, changes in the voter participation rate which may be due to changes in GDP per capita growth does not posit a concern for our estimation strategy.

These graphs show that there is substantial variability in the vote share of extreme political parties, both across time as well as across countries in a given time period. For example, while the average vote share of right-wing/nationalist parties in Denmark was around 8 percent in the 70s, 3 percent in the 80s, and 4 percent in the 90s, in West-Germany the vote share of right-wing/nationalist parties was around 0.3 percent in the 70s, 0.9 percent in the 80s and 2.5 percent in the 90s. In Italy the vote share of right-wing/nationalist parties was around 4 percent in the 70s, 3 percent in the 80s, and 7 percent in the 90s; in France it was around 0 percent in the 70s, 2 percent in the 80s, and 4 percent in the 90s. For the communist parties, the share of votes obtained in Denmark was around 6 percent in the 70s, 11 percent in the 80s, and 10 percent in the 90s. In West-Germany the share of votes obtained by communist parties was around 0.5 percent in the 70s, 0.3 percent in the 80s and 0.5 percent in the 90s. In Italy the share of votes obtained by communist parties was around 14 percent in the 70s, 16 percent in the 80s, and 14 percent in the 90s; and in France it was around 8 percent in the 70s, 6 percent in the 80s, and 5 percent in the 90s.

4 Estimation Strategy

We use the following econometric model to estimate the effect that real per capita GDP growth has on the vote share of extreme political parties:

$$Voteshare_{c,t} = a_c + b_t + cGrowth_{c,t-1} + u_{c,t},$$

where a_c and b_t are country and time fixed effects that capture country-specific unobservables and time-specific common shocks respectively. $u_{c,t}$ is an error term that is clustered at the country level to allow for arbitrary within-country serial correlation. As a baseline regression we use least-squares to estimate the effect that (lagged) real per capita GDP growth has on the vote share of extreme political parties. Note that for our least-squares estimator to provide a consistent estimate of the effect that lagged per capita GDP growth has on the vote share of extreme political parties it is necessary that real per capita GDP growth does not systematically respond to future changes in the share of votes obtained by extreme political parties. Stated differently, this assumption boils down to current investment and labor market decisions being independent of future, predictable changes in the political system. This may be a rather strong assumption that we address econometrically in two ways.

First, we consider using system-GMM estimation (Blundell and Bond, 1998) to estimate a dynamic version of the above equation that uses the lagged vote share as a right-hand-side regressor. Including the lagged vote share on the right-hand side implies that the residual variation in the vote share which correlates with per capita GDP growth is not predictable by agents that use past vote shares to forecast future vote shares. Hence, changes in the current vote share are surprise changes that cannot be predicted by past vote shares. Because these surprise changes cannot be systematically predicted by past vote shares they are less likely to systematically affect past per capita GDP growth due to anticipation effects.

As a second approach to deal with endogeneity issues, we consider instrumental variable techniques that use international oil price shocks as an instrument for real per capita GDP growth. Because the effects of international oil price shocks on real per capita GDP growth are dependent on whether a country is an oil importer or an oil exporter, we construct a country-specific oil price shock series as $Oilshock_{c,t} = \triangle Log(OilPrice_t) * \theta_c$, where $\triangle Log(Oilprice_t)$ is the log-change of the international oil price (obtained from IMF statistics) and θ_c is the country-specific average share of (net) oil exports in GDP (obtained from OECD statistics). Note that we explicitly use a time-invariant net export share to ensure that our oil price shock variable reflects only time-specific movements in the international oil price and not time-specific movements in countries' export-shares. For our oil price shock variable to be a valid instrument we therefore need that country-specific (future) changes in the vote share of extreme political parties do not systematically affect (current) changes in the international oil price. This condition will be satisfied as long as output growth in each OECD country does not significantly affect changes in the international oil price. Or stated differently, that each of our 16 OECD countries is a price taker on the international oil market. According to the International Energy Agency none of our countries has an export or import share that exceeds 5% of total world oil production so changes in the demand or supply of oil to the international oil market which are due to changes in the vote share of extreme political parties in a specific OECD country are likely to have only a negligible effect on the international oil price.

5 Main Empirical Results

Table 2 presents our baseline estimates of the effect that real per capita GDP growth has on the vote share of right-wing/nationalist parties. In column (1) we show the estimates of a least squares regression that does not control for country or time fixed effects. The obtained coefficient on per capita GDP growth in this pooled least-squares regression is negative (-0.071) and statistically significant at the 1 percent level. In column (2) we add the level of per capita GDP to the right-hand-side of the estimating equation. In line with our theoretical predictions from Section 2, the corresponding coefficient on GDP per capita is not significantly different from zero. Moreover, the real per capita GDP growth rate continues to have a highly significant negative effect on the support for extreme right-wing/nationalist parties.

In column (3) we add country fixed effects to account for potential unobservable cross-country heterogeneity. This leaves our point estimate on real per capita GDP growth mostly unchanged. Controlling in column (4) in addition to the country fixed effects for also time fixed effects which capture unobservable shocks common across OECD countries does however make our point estimate increase in absolute size substantially. The point estimate is -0.136 and statistically significant at the 1 percent level. Economically, the estimate implies that a one percentage point decrease in real per capita GDP growth of the prior two quarters increases the vote share of extreme right-wing/nationalist parties in the following period by about 0.136 percentage points.

As an identification check we run in column (5) a false experiment that includes future per capita GDP growth conditional on past per capita GDP growth in the estimating equation. A significant point estimate on future per capita GDP growth could indicate endogeneity problems as a past change in the vote share could affect current GDP per capita growth. We find however that future per capita GDP growth conditional on past per capita GDP growth does not enter the estimating equation with a statistically significant sign and that quantitatively the point estimate on future per capita GDP growth is rather small. Moreover, we find that lagged per capita GDP growth continues to have a statistically significant negative effect on the vote share. In column (6) we also document that per capita GDP growth shocks averaged over the past two years have a significant negative effect on the vote share, pointing towards persistence in the effects that past GDP per capita growth shocks have on current voting behavior.

To check whether our linear specifications miss out on important non-linearities in the relationship between real per capita GDP growth and the vote share of extreme right-wing/nationalist parties we show in Figure 5 non-parametric local polynomial estimates. The nonparametric local polynomial estimates allow for a flexible functional relationship between real per capita GDP growth and the vote share of extreme right-wing/nationalist parties. The estimates are computed using an Epanechnikov kernel, with bandwidth selection based on cross-validation criteria. As can be seen, there is a clear downward sloping relationship between real per capita GDP growth and the vote share of extreme right-wing/nationalist parties over the entire range of real per capita GDP growth. Moreover, the 95% confidence bands indicate that the linear relationship implicitly assumed in our estimating equation cannot be rejected.

In Table 3 we present system-GMM estimates that take into account dynamics in the vote share of extreme right-wing/nationalist parties. The estimated AR(1) coefficient on the vote share is 0.66 and this indicates quite persistent dynamics in our dependent variable. Column (1) also shows that the point estimate on lagged per capita GDP growth in the dynamic panel regression is negative and statistically significant just like in the static panel regression. Note however that the interpretation of the point estimate on the lagged per capita GDP growth variable is slightly different in the dynamic panel regression from the interpretation of the point estimate on the per capita GDP growth variable in the static panel regression because (residual) changes in the vote share are in the dynamic panel regression surprise changes that cannot be forecasted by past changes in the vote share.[10] As column (1) shows, the estimated coefficient on lagged per capita GDP growth is -0.062 and has a t-value of -2.46. The point estimate therefore implies that a permanent decrease in the growth rate of 1 percentage point increases the vote share of extreme right-wing/nationalist parties by over 0.18 percentage points in the long-run. On the other hand, a purely transitory growth shock increases the vote share of extreme right-wing/nationalist parties by 0.06 percentage points on impact and then slowly converges towards zero over time. In column (2) we repeat the exercise using the average real per capita GDP growth rate over the past two years and find similarly to the static panel estimates that past growth shocks have a significant negative effect on the vote share.

[10] Higher order lags of the vote share are not statistically significant.

In Table 4 we further address the issue of possible endogeneity bias in our estimating equation by using international oil price shocks as instrumental variables. The two-stage least squares estimate in column (1) produces a point estimate on lagged per capita GDP growth of -0.998 that is statistically significant at the 1% level. Despite being quantitatively larger than the corresponding least-squares estimate of column (3) in Table 2, a formal Hausman test does not reject the hypothesis that the least-squares estimate is equal to the IV estimate. The first stage F-statistic for the two-stage least squares estimate is around 11.9 so that the maximum relative IV bias is less than 10% according to the tabulations in Stock and Yogo (2005). Moreover, the Hansen J-test does not reject the validity of past oil price growth shocks as instrumental variables for per capita GDP growth. In column (2) we also compute the two-stage least squares estimate for the average real per capita GDP growth rate over the past two years. The first stage F-statistic for this two-stage least squares regression is about 30 and hence easily exceeds the critical values for weak instruments. In the second stage, we obtain a point estimate on lagged per capita GDP growth of -0.374 that is statistically significant at the 5% level. Again we tested the validity of our instruments and did not find evidence that they are systematically correlated with the second stage error.

In Table 5 we report estimates of the effect that economic growth has on the support for communist parties. Our model predicts that the growth effects depend on the stability of redistributive measures over time that voters associate with different parties. We find that the two-stage least squares estimates, reported in columns (1) and (2) of Table 5 do not yield a significant effect of economic growth on the vote share obtained by communist parties.[11] According to our theory, an explanation for this differential effect could be that voters perceive communist parties as being more clearly in favour of redistribution along conventional lines – i.e., from rich to poor – than right-wing/nationalist parties. Note that while the Hausman test does not reject that the least squares estimates reported in columns (3) and (4) of Table 5 are significantly different from the instrumental variables estimates reported in columns (1) and (2) the least squares estimate in column (3) is barely significant at the 10% level and that the least squares estimate in column (4) is not significant at any conventional confidence level.

[11] The corresponding system-GMM estimates, not reported here for space purposes, are also insignificant.

In Table 6 we report the overall effect that economic growth has on the support for extreme political platforms. The two-stage least squares estimate in column (1) of the effect that economic growth has on the combined variables of right-wing/nationalist and communist vote shares is -0.548 (significant at the 1% level). This estimate implies that a decline of growth by three percent would, on average lead to an increase of the vote share of extreme political parties of at most two percentage points. Column (2) shows that the two-stage least squares estimate of the effect that economic growth has on the support for extreme political platforms is also negative when using the real per capita GDP growth rate averaged over the past two years but the t-value in this case is only -1.04 and hence not significant. On the other hand, the respective least squares estimates reported in columns (3)-(4) of Table 6 are both negative and statistically significant at the 1 percent level at least.

According to our theoretical analysis in Section 2, more income inequality should be associated with a smaller effect of growth on the support for extreme political parties. In Table 7 we test for the impact of inequality on the marginal effect that economic growth has on the support for extreme political parties by ordering the countries in our data set according to their median-to-mean income ratio (net of taxes and transfers) and then splitting them into two subsamples with an equal number (8) of countries. Panel A of Table 7 reports the least squares and instrumental variables estimates for the sample with the highest median-to-mean income ratio; Panel B reports the estimates for the sample with the lowest median-to-mean income ratio.[12] As can be seen, the effect of GDP per capita growth on the support for extreme political parties is quantitatively larger and statistically stronger in the group of countries with high median-to-mean income ratios (low inequality) than in the group with low median-to-mean income ratios (high inequality). This result is consistent with our theoretical prediction of a dampening effect of greater income inequality on the marginal effect that economic growth has on the vote share of parties with extreme platforms.[13]

[12]The median median-to-mean after-tax income ratio in Panel A is 0.92; in Panel B the median median-to-mean after-tax income ratio is 0.83. The datasource is OECD (2009) statistics.

[13]Applying the sample split to the right-wing/nationalist parties yields a 2SLS coefficient on economic growth for the low inequality countries of -1.57 that is significant at the 1 percent level, and a 2SLS coefficient for the high inequality countries of -0.04 that is insignificant. For the communist parties the 2SLS estimates are insignificant and quan-

6 Conclusion

Many observers believe that the standard of living and the distribution of income are major determinants of the political support for radical political platforms. Our empirical analysis suggests that economic growth is an important and independent determinant of political radicalism: a lower growth rate increases the support for extreme political platforms.

There are good reasons to believe that industrialized countries' per capita GDP is likely to grow less strongly in the coming decades. Demographic developments impose limits on GDP per capita growth and increasing prices for raw materials make the production in those countries more expensive. The current financial crisis has led to the largest drop in per capita GDP of industrialized countries since the 1930s and the necessity to reduce levels of public debt and the so called for regulation of the financial system may have long-lasting adverse effects on real per capita GDP growth. If Benjamin Friedman is right with his hypothesis, political outcomes could be affected significantly in those economies (see also Miegel, 2009).

The empirical results in this paper instead show that it is unlikely that even strong recessions can change political outcomes. Even a significant drop of the GDP per capita growth rate of three percentage points would increase on average the vote share of the extreme parties considered in our sample by less than two percentage points. Such an increase in the vote share will most likely not change the political outcome in any of the OECD economies substantially.

Our present analysis may be extended into several directions. As the data become available, it is desirable to extend the empirical analysis to developing countries and to other historical episodes.[14] All of the OECD countries in our sample are democracies with a strong historical record of

titatively small regardless of whether we consider the high inequality sample or the low inequality sample.

[14] In the Appendix we have made an attempt to include developing countries in our empirical analysis by using data provided by the Database of Political Institutions (Beck et al., 2001) on the number of seats received by the 1st, 2nd, and 3rd largest party in parliament. The Database of Political Institutions codes whether the 1st, 2nd, and 3rd largest party in parliament has nationalist origin, but unfortunately does not provide information on the number of seats obtained by more minor parties (to which nationalist parties often belong). The Appendix discusses further the pros and cons of using the Database of Political Institutions for our empirical purposes and also presents estimation results.

democracy, and it would be interesting to see whether results also hold in countries that have had little to no experience with democracy. Moreover, our theoretical analysis points out that there may be other determinants of political extremism that should be studied empirically as well as the data become available. One may also extend the theoretical framework to permit different growth effects of different political regimes. In particular, adaptive expectations about growth rates may lead to an interesting dynamic relationship between growth and the political regime. Multiple equilibria may obtain when extreme political regimes grow little which makes individuals believe, that redistribution through the continuation of an extreme political regime is the best way to secure a high living standard.

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Appendix. Results Using the Database of Political Institutions

In this appendix we discuss the use and estimation results for an alternative dataset: the Database of Political Institutions (Beck et al., 2001). The Database of Political Institutions provides information on the number of seats obtained by the 1st, 2nd, and 3rd largest political party voted into parliament and records whether political parties are nationalist. A key requirement for coding a party as nationalist in the Database of Political Institutions is that a primary component of the party's platform is the creation or defense of a national or ethnic identity. According to the Database of Political Institutions examples that fall into the "nationalist" category are parties that have fought for independence, either militarily or politically; parties that advocate the persecution of minorities; or parties that are listed as xenophobic. One clear advantage of the Database of Political Institutions is that it covers a much larger sample of countries than our OECD dataset (180 countries for the period 1975-2006). However, a major disadvantage of the Database of Political Institutions for the purpose of our empirical analysis is that the Database of Political Institutions only provides information on the number of seats obtained by the 1st, 2nd, and 3rd largest political party in parliament. In particular, the database does not provide information on the actual vote share obtained by nationalist parties. For many countries extreme right-wing/nationalist parties only receive a small share of the total number of votes and are therefore not represented as the 1st, 2nd, or 3rd largest party in parliament.

To show what happens when using the information provided by the Database of Political Institutions we present in Appendix Table 1 system-GMM estimates where our dependent variable is the number of seats obtained by a nationalist party in parliament (given that the nationalist party is the 1st (alternatively, 2nd or 3rd) largest party in parliament).[15] Overall we find that there are significant negative effects of past real per capita GDP growth (real per capita GDP growth data are from the Penn World Tables, version 6.3, Heston et al. 2009) on the number of seats received by a nationalist party if the nationalist party constitutes the 2nd or 3rd largest party in

[15]Note that the variable is 0 if the 1st (alternatively, 2nd or 3rd) largest party in parliament is not a nationalist party. The variable is missing if no information was provided on the number of seats received.

parliament (see columns (2) and (3)). For the case of the nationalist party already constituting the largest (i.e. ruling) party in parliament we do not find a statistically significant effect of past per capita GDP growth on the number of seats that the nationalist party received (see column (1)). These results are consistent with the results that we obtained from the survey based vote shares of radical parties in our OECD dataset. Nevertheless, we believe that for purposes of examining empirically how economic growth affects the support for extreme political parties the survey based vote shares are more suitable than the information provided by the Database of Political Institutions on the number of seats obtained by the 1st, 2nd, and 3rd largest party in parliament.

Table 1. Summary Statistics

	Mean	Std. Dev.	Obs.
Vote Share of Right-Wing / Nationalist Parties (Eurobarometer)	0.016	0.027	610
Vote Share of Communist Parties (Eurobarometer)	0.041	0.052	610
Share of Net Oil Exports in GDP (OECD Statistics)	-0.013	0.033	610
GDP Per Capita Growth (OECD Statistics)	0.012	0.054	610
Mean-to-Median After Tax Income Ratio (OECD Statistics)	1.145	0.062	610
GDP Per Capita (OECD Statistics)	16588	11462	610

Table 2. GDP Growth and the Rise of Right-Wing/Nationalist Parties

	(1)	(2)	(3)	(4)	(5)	(6)
	LS	LS	LS	LS	LS	LS
GDP Growth, t-1	-0.071***	-0.073***	-0.071***	-0.136***	-0.126***	
	(-2.73)	(-3.57)	(-2.69)	(-2.80)	(-2.93)	
GDP Level, t-1		-6.60×10^{-8}				
		(-0.17)				
GDP Growth, t+1					-0.030	
					(-0.84)	
Average GDP Growth, t-1 to t-4						-0.076**
						(-2.34)
Within-Country R^2	0.086	0.086	0.086	0.156	0.156	0.186
Country FE	No	No	Yes	Yes	Yes	Yes
Year FE	No	No	No	Yes	Yes	Yes
Observations	610	610	610	610	610	610

Note: The method of estimation is least squares. The t-values listed in parentheses are based on Huber robust standard errors that are clustered at the country level. The dependent variable is the share of survey votes received by right-wing / nationalist parties. *Significantly different from zero at 90 percent confidence, ** 95 percent confidence, *** 99 percent confidence.

Table 3. GDP Growth and the Rise of Right-Wing/Nationalist Parties

	(1)	(2)
	SYS-GMM	SYS-GMM
Voteshare, t-1	0.657***	0.779***
	(11.88)	(20.40)
GDP Growth, t-1	-0.062**	
	(-2.46)	
Average GDP Growth, t-1 to t-4		-0.025***
		(-2.01)
AR (2) Test, p-value	0.122	0.106
Sargan Test, p-value	0.166	0.127
Country FE	Yes	Yes
Year FE	Yes	Yes
Observations	530	530

Note: The method of estimation is system-GMM (Blundell and Bond, 1998). The t-values shown in parentheses are based on Huber robust standard errors that are clustered at the country level. The dependent variable is the share of survey votes received by right-wing / nationalist parties. *Significantly different from zero at 90 percent confidence, ** 95 percent confidence, *** 99 percent confidence.

Table 4. GDP Growth and the Rise of Right-Wing/Nationalist Parties

	(1)	(2)
	2SLS	2SLS
GDP Growth, t-1	-0.998***	
	(-3.69)	
Average GDP Growth, t-1 to t-4		-0.374**
		(-2.42)
First Stage F-stat	11.898	30.938
Hansen Overid. Test, p-value	0.2145	0.8597
Hausman Endogeneity Test, p-value	0.5802	0.4902
Country FE	Yes	Yes
Year FE	Yes	Yes
Observations	610	610

Note: The method of estimation is two-stage least squares. The t-values shown in parentheses are based on Huber robust standard errors that are clustered at the country level. The instrumental variables are the *t*-2 to *t*-4 oil price growth rate weighted by the country-specific (time-invariant) net export share of oil in GDP (see Section 4 for a detailed description of the instrument). The dependent variable is the share of survey votes received by right-wing / nationalist parties. *Significantly different from zero at 90 percent confidence, ** 95 percent confidence, *** 99 percent confidence.

Table 5. GDP Growth and the Rise of Communist Parties

	(1)	(2)	(3)	(4)
	2SLS	2SLS	LS	LS
GDP Growth, t-1	0.443		-0.102*	
	(1.57)		(-1.69)	
Average GDP Growth, t-1 to t-4		0.250		-0.035
		(1.23)		(-1.04)
First Stage F-stat	11.898	30.938	.	.
Hansen Overid. Test, p-value	0.353	0.732	.	.
Hausman Endogeneity Test, p-value	0.334	0.584	.	.
Country FE	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes
Observations	610	610	610	610

Note: The method of estimation in columns (1) and (2) is two-stage least squares; columns (3) and (4) least squares. The t-values shown in parentheses are based on Huber robust standard errors that are clustered at the country level. The instrumental variables are the *t*-2 to *t*-4 oil price growth rate weighted by the country-specific (time-invariant) net export share of oil in GDP (see Section 4 for a detailed description of the instrument). The dependent variable is the share of survey votes received by communist parties. *Significantly different from zero at 90 percent confidence, ** 95 percent confidence, *** 99 percent confidence.

Table 6. GDP Growth and the Rise of Extreme Political Parties

	(1)	(2)	(3)	(4)
	2SLS	2SLS	LS	LS
GDP Growth, t-1	-0.548***		-0.239***	
	(-3.34)		(-2.97)	
Average GDP Growth, t-1 to t-4		-0.125		-0.111***
		(-1.04)		(-2.67)
First Stage F-stat	11.898	30.938	.	.
Hansen Overid. Test, p-value	0.479	0.811	.	.
Hausman Endogeneity Test, p-value	0.850	0.536	.	.
Country FE	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes
Observations	610	610	610	610

Note: The method of estimation in columns (1) and (2) is two-stage least squares; columns (3) and (4) least squares. The t-values shown in parentheses are based on Huber robust standard errors that are clustered at the country level. The instrumental variables are the *t*-2 to *t*-4 oil price growth rate weighted by the country-specific (time-invariant) net export share of oil in GDP (see Section 4 for a detailed description of the instrument). The dependent variable is the sum of the share of survey votes received by right-wing/nationalist parties and communist parties. *Significantly different from zero at 90 percent confidence, ** 95 percent confidence, *** 99 percent confidence.

Table 7. GDP Growth, Inequality, and the Rise of Political Extremism

Panel A: 8 OECD Countries With Highest Median to Mean Income Ratio (Low Inequality)

	(1)	(2)	(3)	(4)
	2SLS	2SLS	LS	LS
GDP Growth, t-1	-0.678**		-0.229***	
	(-2.40)		(-3.13)	
Average GDP Growth, t-1 to t-4		-0.124**		-0.090**
		(-2.02)		(-2.08)
First Stage F-stat	63.704	61.059	.	.
Hansen Overid. Test, p-value	0.8895	0.5257	.	.
Hausman Endogeneity Test, p-value	0.3750	0.9191	.	.
Country FE	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes
Observations	253	253	253	253

Panel B: 8 OECD Countries With Lowest Median-to-Mean Income Ratio (High Inequality)

	(1)	(2)	(3)	(4)
	2SLS	2SLS	LS	LS
GDP Growth, t-1	-0.280		-0.167	
	(-0.75)		(-1.47)	
Average GDP Growth, t-1 to t-4		-0.004		-0.084
		(-0.06)		(-1.55)
First Stage F-stat	9.705	2.918		
Hansen Overid. Test, p-value	0.707	0.523		
Hausman Endogeneity Test, p-value	0.290	0.242		
Country FE	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes
Observations	357	357	357	357

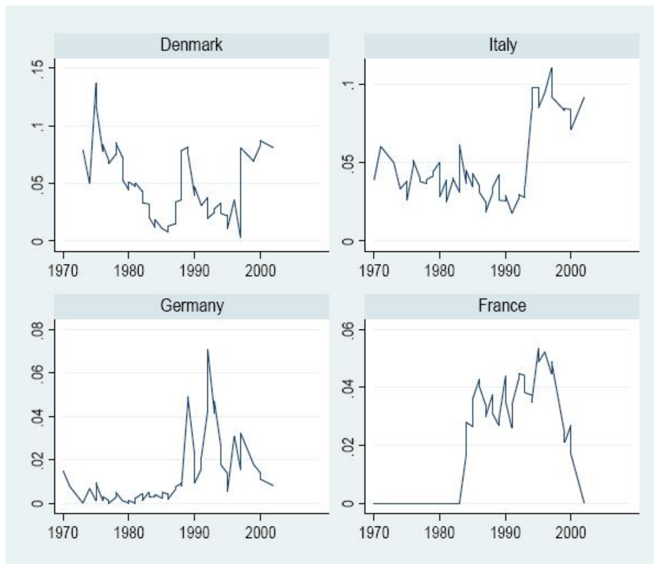
Note: The method of estimation in columns (1) and (2) is two-stage least squares; columns (3) and (4) least squares. The t-values listed in parentheses are based on Huber robust standard errors that are clustered at the country level. The instrumental variables for the two-stage least squares estimation are the *t*-2 to *t*-4 oil price growth rate weighted by the country-specific (time-invariant) net export share of oil in GDP (see Section 4 for a detailed description of the instruments). The dependent variable is the share of survey votes received by right-wing / nationalist parties and communist parties. *Significantly different from zero at 90 percent confidence, ** 95 percent confidence, *** 99 percent confidence.

Appendix Table 1. GDP Growth and the Rise of Political Extremism
(Database of Political Institutions)

	1st Largest Party	2nd Largest Party	3rd Largest Party
	(1)	(2)	(3)
	SYS-GMM	SYS-GMM	SYS-GMM
GDP Growth, t-1	-8.418	-10.181	-7.100
	(-0.31)	(-0.99)	(-0.68)
GDP Growth, t-2	-1.167	-21.395**	-14.613*
	(-0.02)	(-2.18)	(-1.85)
GDP Growth, t-3	-26.161	-19.585	-0.246
	(-1.41)	(-1.59)	(-0.03)
GDP Growth, t-4	62.808	-27.158*	-7.034**
	(0.98)	(-1.69)	(-2.06)
GDP Growth, t-5	-52.655	-13.947	-10.359
	(-1.09)	(-1.55)	(-0.85)
Number of Seats in Parliament, t-1	0.920***	0.716***	0.796***
	(26.86)	(6.21)	(44.69)
Country FE	Yes	Yes	Yes
Year FE	Yes	Yes	Yes
Observations	3466	1293	742

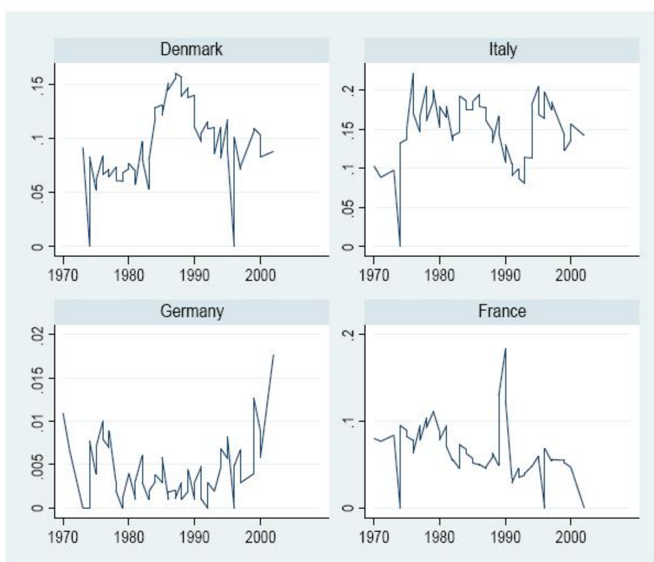
Note: The method of estimation is system-GMM (Blundell and Bond, 1998). The t-values listed in parentheses are based on Huber robust standard errors that are clustered at the country level. The dependent variable in column (1) is the number of seats in parliament that are obtained by a nationalist party if the party constitutes the 1st largest party in parliament; column (2) the number of seats in parliament that are obtained by a nationalist party if the party constitutes the 2nd largest party in parliament; the number of seats in parliament that are obtained by a nationalist party if the party constitutes the 3rd largest party in parliament. *Significantly different from zero at 90 percent confidence, ** 95 percent confidence, *** 99 percent confidence.

Figure 1. Time-Series Plots of the Vote Shares of Right-Wing / Nationalist Parties



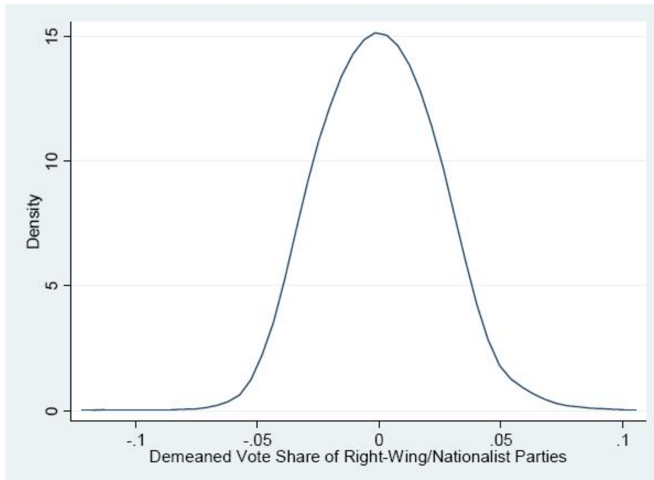
Source: Eurobarometer. The figure is based on answers to the question: "If there were general elections tomorrow, which party would you vote for". Right-wing/nationalist parties are classified according to the ZEUS party code.

Figure 2. Time-Series Plots of the Vote Shares of Communist Parties



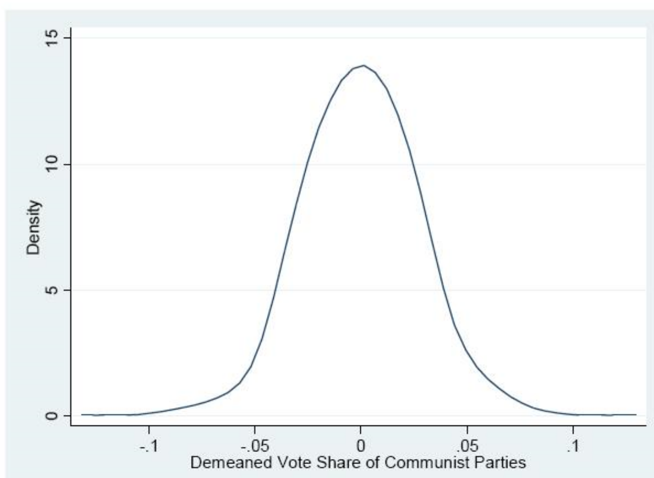
Source: Eurobarometer. The figure is based on answers to the question: "If there were general elections tomorrow, which party would you vote for". Communist parties are classified according to the ZEUS party code.

Figure 3. Kernel Density Plot of Demeaned Vote Shares of Right-Wing / Nationalist Parties



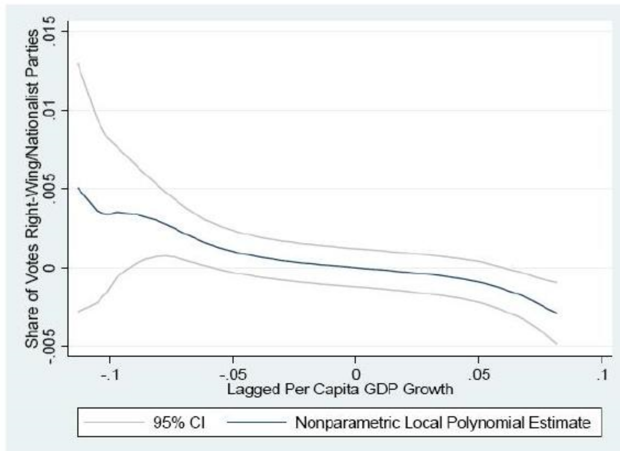
Note: The figure shows an Epanechnikov kernel density estimate of the vote share of right-wing / nationalist parties. The vote share has been demeaned from the country and time fixed effect.

Figure 4. Kernel Density Plot of Demeaned Vote Shares of Communist Parties



Note: The figure shows an Epanechnikov kernel density estimate of the vote share of communist parties. The vote share has been demeaned from the country and time fixed effect.

Figure 5. Per Capita GDP Growth and Vote Shares of Right-Wing / Nationalist Parties



Note: The figure shows nonparametric local polynomial estimates (based on an Epanechnikov kernel) of the relationship between the share of votes obtained by right-wing / nationalist parties and lagged per capita GDP growth. Both the share of votes obtained by right-wing / nationalist parties as well as lagged per capita GDP growth have been demeaned from the country and time fixed effect.