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Does austerity pay off?

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Does austerity pay off?*

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Abstract

Austerity measures are frequently enacted when the sustainability of public finances is in doubt. Such doubts are reflected in high sovereign yield spreads and put further strain on government finances. Is austerity successful in restoring market confidence, bringing about a reduction in yield spreads? We employ a new panel data set which contains sovereign yield spreads for 26 emerging and advanced economies and estimate the effects of cuts of government consumption on yield spreads and economic activity. The conditions under which austerity takes place are crucial. During times of fiscal stress, spreads rise in response to the spending cuts, at least in the short-run. In contrast, austerity pays off, if conditions are more benign.

Keywords: Fiscal Policy, austerity, sovereign risk, yield spreads, confidence, panel VAR, local projections

JEL-Codes: E62

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“Debt is a slow-moving variable that cannot—and in general should not—be brought down too quickly. But interest rates can change much more quickly than fiscal policy and debt.” (C. Reinhart and K. Rogoff 2013)

“[F]inancial investors are schizophrenic about fiscal consolidation and growth. They react positively to news of fiscal consolidation, but then react negatively later, when consolidation leads to lower growth—which it often does.” (O. Blanchard 2011)

1 Introduction

In the years following the global financial crisis, many European governments have been implementing sizeable austerity measures: spending cuts and tax increases in order to reduce budget deficits. These measures were taken in order to confront mounting concerns about rising levels of public debt or outright solvency issues. In fact, in a number of euro area countries sovereign yield spreads relative to Germany started to take off by 2010, arguably leaving policy makers with no alternative course of action.¹ Yet a further deterioration of financial market conditions, coupled with dismal growth performance in the following years, lead many observers to question the wisdom of austerity. In this paper, we ask whether austerity pays off, that is, whether it actually helps restoring market confidence in the sustainability of public finances as reflected by sovereign yield spreads.

We focus on how financial markets respond to austerity measures or, more specifically, on the response of yield spreads and sidestep the issue of how such measures impact the actual health of government finances. In fact, while the response of fiscal indicators such as the level of sovereign debt is of first-order importance in this regard, it generally does not provide a sufficient statistic for assessing the sustainability of debt. For the willingness and the ability of governments to honor a given level of debt obligations depends on a number of country-specific, partly unobserved factors, such as the ability to raise taxes. The same level of debt may thus have very different implications for debt sustainability in different countries (Bi, 2012). Sovereign yield spreads, instead, provide more comprehensive picture, both because they reflect a broader assessment of market participants and because

¹Historically, in addition to primary surpluses, output growth as well as negative real interest rates have contributed to the reduction of debt-to-GDP ratios (Hall and Sargent, 2011). Real interest rates in turn may have been depressed due to “financial repression” (Reinhart and Sbrancia, 2014). While it is unclear to what extent these factors will play an important role in stabilizing debt levels in the years to come, they are arguably no viable means in order to meet market pressures instantaneously.

of the immediate budgetary consequences of higher interest rates (see, e.g., Lorenzoni and Werning, 2014).

Among the many factors which enter a comprehensive assessment, output growth or, more generally, the level of economic activity, plays a key role, because it determines the amount of resources available for debt service (see, e.g., Arellano, 2008). In addition to debt levels and deficits, the growth performance of countries is therefore closely monitored by financial market participants. Depending on how austerity measures impact output growth, yield spreads may thus either rise or fall in response to austerity. Some observers indeed suggest that financial markets are “schizophrenic” about austerity in that they demand austerity measures as public debt builds up, but fail to reward them, as austerity slows down output growth (Blanchard, 2011; Cotarelli and Jaramillo, 2012).

However, the output effect of austerity measures, captured by the fiscal multiplier, is itself heavily debated. More importantly still, recent contributions have stressed the state dependence of fiscal multipliers, that is, their tendency to change with the economic environment. Given the issue at hand, it is particularly noteworthy that a number of studies suggest that the multiplier is less positive or even negative whenever public debt is high (Corsetti, Meier, and Müller, 2012a; Auerbach and Gorodnichenko, 2013a; Ilzetki, Mendoza, and Végh, 2013).² These results, however, are subject to the caveat that they condition on debt, rather than on a more comprehensive measure of “fiscal stress”. Moreover, the critical value for public debt is specified in an arbitrary manner. In our analysis below, we will account for the possibility that the effects of austerity change with the level of fiscal stress, drawing as a conditioning variable on sovereign yield spreads.

As a first step of our analysis, we set up a new data set for sovereign yield spreads. Specifically, we compute spreads for 26 advanced and emerging countries as the difference in sovereign yields vis-à-vis a “riskless” reference country where sovereign default is considered to be extremely unlikely. Importantly, we only consider yields on government securities issued in a common currency in order to eliminate the confounding effects of inflation and depreciation expectations and to isolate market expectations of an outright default. It turns out that our spread measure moves in close sync with credit default swap spreads

²On the other hand, fiscal measures tend to impact the economy more strongly if there is pervasive slack in the economy (Auerbach and Gorodnichenko, 2012) or if the economy is stuck at the zero lower bound (Christiano, Eichenbaum, and Rebelo, 2011). Depending on the state of the economy, it may thus be beneficial in terms of macroeconomic outcomes to either frontload or to delay austerity measures (Corsetti, Kuester, Meier, and Müller, 2010). More extreme still, hysteresis effects may make austerity measures self-defeating to the extent that contractionary fiscal measures may raise the long-run financing costs of governments (De Long and Summers, 2012).

which are available only for parts of our sample. Moreover, our measure of yield spreads measures the real borrowing costs of governments.

We establish a number of basic facts regarding yield spreads. First, they vary considerably across time and countries. In some instances they are slightly below zero, in others they are as high as 70 percentage points. A total of about 1500 observations also allows us to compute the empirical density function. It increases sharply at low levels of spreads, as the number of observations for which spreads are high is limited. Second, yield spreads co-move negatively with economic activity. The correlation of yield spreads and current output growth is negative in all countries of our sample, but Sweden. Third, there is no systematic correlation pattern of spreads and government consumption.

In a second step, we provide estimates on the effects of austerity. We focus on the effects of government consumption cuts, as identification is somewhat less controversial in this case than in case of tax hikes. Specifically, we assume that government consumption is predetermined within a given quarter. This assumption goes back to Blanchard and Perotti (2002) and is rationalized by the fact that changes in government spending cannot be agreed upon without a considerable decision lag. We collect quarterly data for government expenditure while drawing on earlier work by Ilzetki, Mendoza, and Végh (2013), extending their data set to include observations up to the year 2013. For some countries, our observations for both quarterly government consumption as well as sovereign yield spreads date back to the beginning of the 1990s.

We pursue alternative econometric strategies to obtain estimates for how government consumption impacts the economy and, eventually, sovereign yield spreads. Following a large empirical literature on the fiscal transmission mechanism, we rely on estimated panel vector autoregressions (VAR) to compute impulse response functions. In addition, we also employ local projections which, following Jordá (2005), have become very popular in recent years. For both models we compute unconditional estimates for the effects of fiscal shocks as well as estimates conditional on the state of the economy. In this regard, local projections stand out in terms of flexibility (Auerbach and Gorodnichenko, 2013a).

Our main results can be summarized as follows. We find that austerity, or more precisely, cuts of government consumption, tend to raise sovereign yield spreads, if we do not condition our estimates on fiscal stress. At the same time output declines considerably. Upon closer inspection, however, these results mask considerable heterogeneity. If we condition the effects of spending cuts on the absence of fiscal stress, we find that spreads decline. At the same time, we find fiscal multipliers much reduced. In fact, we cannot reject the hypothesis

that government consumption leaves economic activity unaffected in the absence of fiscal stress. On the other hand, in the presence of fiscal stress spending cuts raise spreads and depress economic activity considerably. Fiscal stress episodes thus tend to dominate the overall sample. We also find that spreads tend to decline in response to spending cuts in the medium to long run.

Our results are based on exogenous variations in government consumption, while austerity is typically considered to be an endogenous response to, say, financial market developments. That said, note that identifying an exogenous variation in government consumption is key to isolate the impact of austerity on the variables of interest *per se*, rather than the joint effect of financial market developments *cum* austerity. Still, it is certainly possible that austerity measures impact the economy in different ways than a “regular” fiscal shock—perhaps because they are implemented under special circumstances. Conditioning the effects of spending cuts on the state of the economy is one strategy to address this concern.

Alternative, and complementary, approaches to assess the effects of fiscal consolidation episodes include case studies, notably the classic analysis of Giavazzi and Pagano (1990). Yet another approach goes back to Alesina and Perotti (1995), recently applied by Alesina and Ardagna (2013). It identifies (large) fiscal adjustments as episodes during which the cyclically adjusted primary deficit falls relative to GDP by a certain amount. Finally, fiscal consolidations have also been identified on the basis of a narrative approach (Guajardo, Leigh, and Pescatori, 2011; Devries, Guajardo, Leigh, and Pescatori, 2011).

Our analysis also provides new results on the effects of government spending cuts on sovereign yield spreads for a large panel of advanced and emerging economies. Related studies include numerous attempts to assess the effects of fiscal policy on interest rates. In particular, Ardagna (2009) finds that interest rates tend to decline in response to large fiscal consolidations. Laubach (2009) investigates how changes in the U.S. fiscal outlook affect interest rates. Finally, Akitoby and Stratmann (2008) use a similar measure for sovereign yield spreads as we use in the present paper. They focus on emerging market economies, however, and assess the contemporaneous impact of fiscal variables on spreads within a given year.

The remainder of the paper is organized as follows. Section 2 details the construction of our data set. In this section, we also establish a number of basic facts regarding the time-series properties of sovereign yield spreads and their relationship to government consumption and output growth. In Section 3 we discuss our econometric specification and identification strategy. We present the main results of the paper and an extensive sensitivity analysis in

Section 4. Section 5 concludes.

2 Data

Our analysis is based on a new data set. It contains quarterly data for government consumption, output, and sovereign yield spreads for 26 emerging and advanced economies. Our econometric strategy below requires the use of quarterly data. While data on yield spreads are available at higher frequency, data on macroeconomic aggregates are not. In fact, for a long time, time-series studies of the fiscal transmission mechanism have been limited to a small set of countries, because data for government consumption has not been available at (non-interpolated) quarterly frequency.³ In a recent contribution, Ilzetki, Mendoza, and Végh (2013) have collected non-interpolated quarterly data for government consumption for 44 countries. We reconstruct quarterly data for government consumption along the lines of Ilzetki, Mendoza, and Végh (2013) for the subset of countries for which we are also able to compute sovereign yield spreads. In the process, we also extend the sample to include more recent observations. Our earliest observations for which we have both spread and government consumption data is 1990Q1 for Denmark and Ecuador. Our sample runs up to 2013.

Table 1 provides summary statistics for the government consumption-to-GDP ratio. Government consumption is exhaustive government consumption and does not include transfer payments or government investment. As in Ilzetki, Mendoza, and Végh (2013), depending on the availability of quarterly time series, it pertains to either the general or the central government. The ratio of government consumption-to-GDP varies both across time and across countries. In case of general government data, government consumption fluctuates around 20 percent of GDP.

As a distinct contribution, we also construct a panel data set for sovereign yield spreads in order to measure the assessment of financial markets regarding the sustainability of public finances. Given observations on quarterly government consumption, we aim to construct measures of yield spreads for as many countries as possible. As stressed in the introduction, we construct yield spreads using yields for securities issued in common currency. To the extent that goods and financial markets are sufficiently integrated, we are thereby

³Some studies have resorted to annual data (e.g. Beetsma, Giuliodori, and Klaasen, 2006, 2008; Bénétrix and Lane, 2013). In this case identification assumptions tend to be more restrictive. However, Born and Müller (2012) consider both quarterly and annual data for four OECD countries. They find that the estimated effects of government spending shocks do hardly differ.

Table 1: Basic properties of government consumption-to-output ratio

Country	first obs	last obs	min	max	mean	std
Belgium	1995.00	2013.25	0.24	0.25	0.25	0.00
Denmark	1991.00	2013.25	0.25	0.29	0.27	0.01
Finland	1990.00	2013.25	0.18	0.26	0.21	0.02
France	1986.00	2013.25	0.23	0.27	0.25	0.01
Greece	2000.00	2011.00	0.17	0.23	0.18	0.01
Hungary	1995.00	2013.25	0.21	0.28	0.23	0.02
Ireland	1997.00	2013.25	0.16	0.18	0.17	0.01
Italy	1991.00	2013.25	0.19	0.22	0.20	0.01
Netherlands	1988.00	2013.25	0.23	0.28	0.25	0.01
Poland	1995.00	2013.25	0.17	0.21	0.18	0.01
Portugal	1995.00	2013.25	0.19	0.22	0.20	0.01
Slovenia	1995.00	2013.25	0.17	0.22	0.19	0.01
Spain	1995.00	2013.25	0.16	0.22	0.18	0.02
Sweden	1993.00	2013.25	0.07	0.12	0.09	0.02
United Kingdom	1986.00	2013.25	0.21	0.28	0.23	0.02
Argentina	1993.00	2013.25	0.12	0.15	0.13	0.01
Chile	1989.00	2012.75	NaN	NaN	NaN	NaN
Colombia	2000.00	2013.25	0.15	0.17	0.16	0.00
Ecuador	2000.00	2013.25	0.10	0.13	0.12	0.01
El Salvador	1994.00	2013.25	0.06	0.09	0.07	0.01
Malaysia	2000.00	2013.00	NaN	NaN	NaN	NaN
Peru	1995.00	2013.25	NaN	NaN	NaN	NaN
South Africa	1993.00	2013.25	0.13	0.18	0.15	0.02
Thailand	1993.00	2013.25	0.07	0.11	0.09	0.01
Turkey	1998.00	2013.25	0.09	0.12	0.11	0.01
Uruguay	1988.00	2013.25	NaN	NaN	NaN	NaN

Notes: Government consumption is consumption of the general government except for Chile, El Salvador, Malaysia, Peru, and Sweden, where it refers to central government consumption. For Chile, Malaysia, Uruguay, and Peru, we do not have information about the *level* of quarterly non-interpolated government consumption.

eliminating fluctuations in yields due to changes in real interest rates, inflation expectations and the risk premia associated with them. In addition to a default risk premium, yield spreads may still reflect liquidity premium and, if duration differs or drifts, a term premium (see e.g. Broner, Lorenzoni, and Schmunkler, 2013). However, we try to minimize the term premium by constructing the yield spread on the basis of yields for bonds with a comparable maturity and coupon.⁴ Moreover, any liquidity premium is likely to be small in

⁴We focus on long-term rates whenever possible. As they are closely linked to the average of expected future short-term rates, they are a more appropriate measure of governments' refinancing costs than short-term-term rates. Assessing the effects of austerity on the term structure is beyond the scope of the present

our sample. As a result, yield spreads should reflect primarily financial markets' assessment of the probability and extent of debt repudiation by a sovereign.

We obtain spreads following three distinct strategies. First, for a subset of (formerly) emerging markets we directly rely on J.P. Morgan's Emerging Market Bond Index (EMBI) spreads which measure the difference in yields of dollar-denominated government or government-guaranteed bonds of a country relative to those of U.S. government bonds.⁵ Second, we add to those observations data for euro area countries based on the "long-term interest rate for convergence purposes". Those are computed as "yields to maturity" according to the International Securities Market Association (ISMA) formula 6.3 from "long-term government bonds or comparable securities" with a residual maturity of close to 10 years (ideally 9.5 to 10.5 years) with a sufficient liquidity (see, for details, European Central Bank, 2004). In case more than one bond is included in the sample, simple averaging over yields is performed to obtain a representative rate. For this country group, we use the German government bond yield as the risk-free benchmark rate and compute spreads relative to the German rate.⁶

Finally, we also make use of the issuance of foreign currency government bonds in many advanced economies during the 1990s and 2000s to extend our sample to non-euro countries and the pre-EMU period. In particular, drawing on earlier work by Bernoth, von Hagen, and Schuknecht (2012), we identify bonds denominated in either US dollar or Deutsche mark of at least 5 years of maturity by developed countries. We compute the spread of yields for those bonds relative to the yields of US or German government bonds of comparable maturity and coupon yield in order to have comparable duration and thus term premia.⁷ Whenever possible, we aim to minimize the difference in coupon yield to 25 basis points and the difference in maturity to one year. In order to avoid artifacts introduced by trading drying up in the last days before redemption, we omit the last thirty trading days before

study.

⁵Note that inclusion of a bond into the EMBI requires a minimum bond issue size of \$500 million, assuring that the liquidity premium compared to US bonds is not too large. Moreover, we rely on stripped spreads (Datastream Mnemonic: SSPRD), which "strip" out collateral and guarantees from the calculation.

⁶The bonds used for computing the "long-term interest rate for convergence purposes" are typically bonds issued in euro, but under national law. In this regard they differ from the securities on which the EMBI is based, which are typically issued under international law. This difference becomes important if the monetary union is believed to be reversible. In case of exit from the EMU, the euro bonds will most likely be converted into domestic currency bonds, implying that they should carry a depreciation/exchange rate premium that is absent in the international law bonds. Still, even during the height of the European debt crisis, reversibility risk accounted for a small fraction of sovereign yield spreads in Greece (Kriwoluzky, Müller, and Wolf, 2014).

⁷Yields on individual bonds are based on the yield to maturity at the midpoint as reported in Bloomberg or the yield to redemption in Datastream.

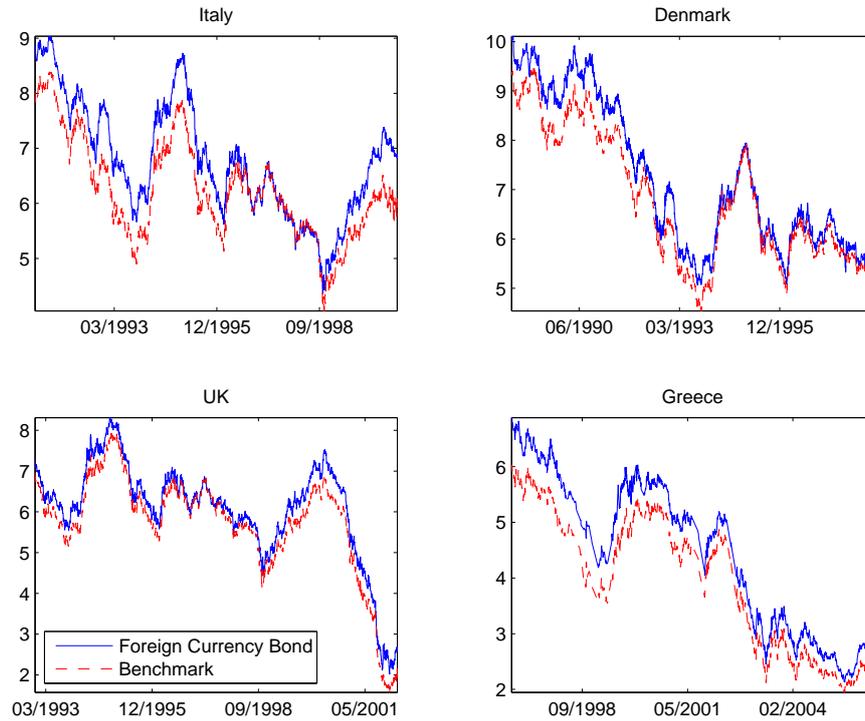


Figure 1: Bond yields to maturity for selected countries. Notes: blue solid line: yield to maturity/redemption yield of the respective domestic bond issued in foreign currency; red dashed line: yield to maturity/redemption yield of a “risk-free” benchmark bond with comparable coupon yield and maturity.

the earliest maturity date of either benchmark or the government bond.⁸ In case of several bonds being available for overlapping periods, we average over yield spreads. In order to construct quarterly time-series data, we average yield spreads over the trading days of a given quarter.⁹ Note also that by focusing on common currency bonds, our spread measure is not affected by the convergence play observed for nominal yield spreads prior to the introduction of the euro.

Figure 1 provides some examples of how sovereign yield spreads are constructed. For

⁸Still, in moving along the yield curve, we may pick up cross-country differences in the slope of the yield curve. In principle, this effect can be quantitatively significant (Broner, Lorenzoni, and Schmunkler, 2013). However, as we find our spread measure to comove very strongly with CDS spreads (whenever they are available), we ignore the issue in the present paper.

⁹Our procedure thus mimics the creation of the EMBI spreads and “long-term interest rate for convergence purposes”. We note, however, that we rely on a smaller foreign currency bond universe and cannot correct for maturity drift (Broner, Lorenzoni, and Schmunkler, 2013). Because of this, we use these data only for the shortest necessary sample, that is, for EMU countries we rely on the “long-term interest rate for convergence purposes” whenever available.

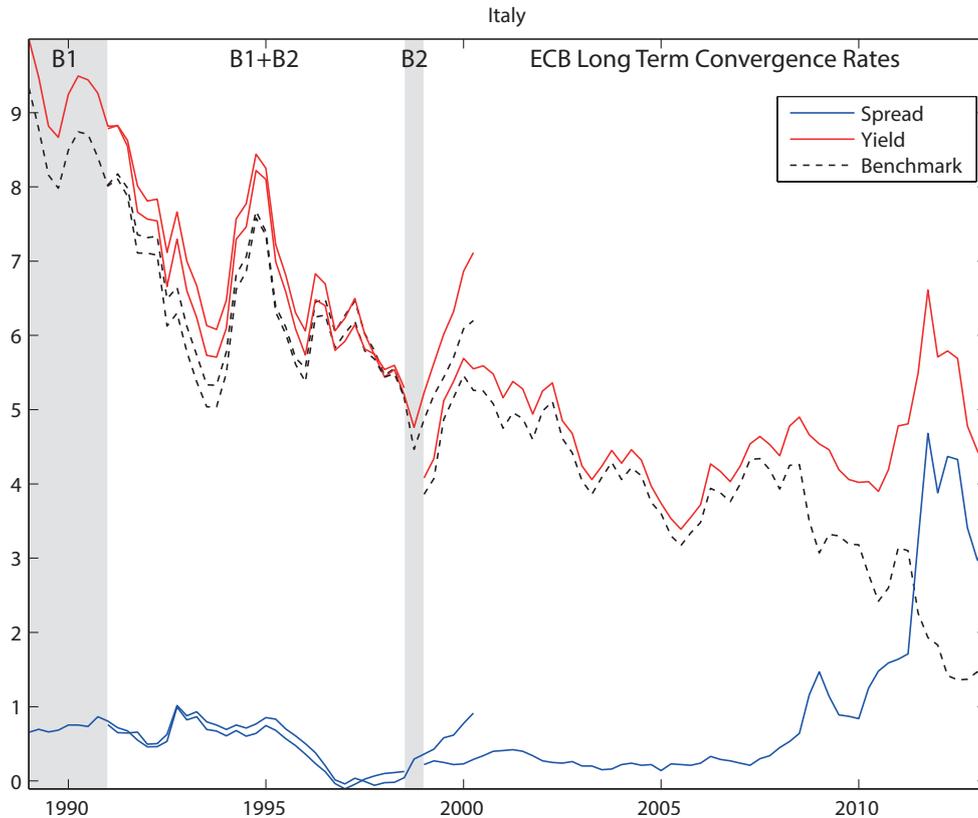


Figure 2: Construction of the Italian yield spread series.

four countries, it displays the yields of foreign currency bonds jointly with those of the associated benchmark bonds. For three countries (Italy, Denmark, UK), we consider bonds denominated in US dollars, while for Greece we consider a bond issued in Deutsche mark.¹⁰ Note that yield spreads are typically small relative to the level of yields and vary considerably over time. For Italy and Greece, data on foreign currency bonds allow us to extend the sample to include observations prior to the introduction of the euro. In case of Denmark and the UK, they allow us to compute common-currency yield spreads, although those countries are not members of the euro zone.

Figure 2 details the construction for the case of Italy. Until 1991, only one foreign bond is

¹⁰Italy: 10year \$US bond issued on 08/02/1991 with coupon 8.75% (XS0030152895); benchmark bond: US 10 year Treasury note issued on 15/11/1990 with coupon 8.25% (US912827ZN50). Denmark: 10year \$US zero coupon bond issued on 8/6/1986 (GB0042654961); benchmark bond: US 10 year Treasury note issued on 15/08/1998 with coupon 9.25% (ISIN: US912827WN87). UK: 10year \$US bond issued on 12/9/1992 with coupon 7.25% (XS0041132845); benchmark bond: US 10 year Treasury note issued on 06/05/1992 with coupon 7.5% (US912827F496). Greece: 10year DEM bond issued on 11/13/1996 with coupon 6.75% (DE0001349355); benchmark bond: German 15 year bond issued on 04/10/1996 with coupon 6.25% (DE0001135010).

available. Starting in 1992, we obtain a second bond and compute the yield spread as the average over those of both bonds. When the first bond matures in 1997, we are left with one bond until 1999. From that point on, we use the long-term convergence bond yields provided by the ECB.

Table 2: Basic properties of sovereign yield spreads

Country	first obs	last obs	min	max	mean	std	$\rho(\Delta y_t, s_t)$	$\rho(\Delta g_t, s_t)$
Belgium	1991.75	2013.25	0.04	2.53	0.45	0.44	-0.38	-0.13
Denmark	1988.50	2002.50	0.02	1.93	0.57	0.42	-0.17	-0.01
Finland	1992.25	2013.25	-0.04	0.80	0.27	0.18	-0.46	-0.28
France	1999.00	2013.25	0.02	1.35	0.27	0.32	-0.33	0.03
Greece	1992.25	2013.25	0.15	23.98	2.80	5.24	-0.59	-0.22
Hungary	1999.00	2013.25	0.10	6.05	1.79	1.66	-0.63	-0.05
Ireland	1991.75	2013.25	-0.04	7.92	1.04	1.79	-0.23	-0.43
Italy	1989.00	2013.25	-0.07	4.68	0.77	0.98	-0.41	-0.42
Netherlands	1999.00	2013.25	0.00	0.67	0.19	0.17	-0.63	-0.15
Poland	1994.75	2013.25	0.42	8.71	1.97	1.43	-0.01	-0.11
Portugal	1993.25	2013.25	0.00	11.39	1.27	2.62	-0.44	-0.41
Slovenia	2006.00	2013.25	-0.17	5.13	1.59	1.59	-0.42	-0.42
Spain	1992.50	2013.25	0.01	5.07	0.71	1.15	-0.61	-0.50
Sweden	1986.00	2009.50	-0.95	2.95	0.90	0.94	0.33	-0.09
United Kingdom	1992.75	2007.75	-0.03	0.64	0.29	0.17	-0.18	0.06
Argentina	1993.75	2013.25	2.04	70.78	15.80	18.74	-0.09	-0.21
Chile	1999.25	2013.25	0.55	3.43	1.45	0.58	-0.45	0.01
Colombia	1997.00	2013.25	1.12	10.66	3.65	2.19	-0.38	-0.19
Ecuador	1995.00	2013.25	5.02	47.64	12.58	9.08	-0.12	0.03
El Salvador	2002.25	2013.25	1.27	8.54	3.33	1.36	-0.47	0.04
Malaysia	1996.75	2013.25	0.46	10.55	1.84	1.45	-0.62	0.03
Peru	1997.00	2013.25	1.14	9.11	3.60	2.01	-0.36	-0.05
South Africa	1994.75	2013.25	0.70	6.52	2.28	1.24	-0.43	-0.34
Thailand	1997.25	2006.00	0.48	5.55	1.51	1.11	-0.55	0.15
Turkey	1996.25	2013.25	1.39	10.66	4.19	2.44	-0.31	-0.15
Uruguay	2001.25	2013.25	1.27	16.43	4.07	3.18	-0.36	-0.29

Notes: spreads s_t are average of daily observations per quarter, measured in percentage points.

Table 2 provides information on the coverage of our spread sample and some basic descriptive statistics.¹¹ Spreads s_t are measured in percentage points and vary considerably across

¹¹Figure A.1 in the appendix displays the time series on a country-by-country basis.

our sample. In a couple of countries the lowest realization of the spread is negative, that is, yields fall below those of the reference bond. For the advanced economies group¹² we observe the highest spreads in Portugal (11) and Greece (24). For the emerging economies¹³ the highest values are reached in Ecuador (48) and Argentina (71).

Measured relative to these values, most realizations of spreads in our sample are small. This is apparent from the empirical distribution function (CDF) which we plot in Figure 3 for the entire sample (solid line), but also for the set of advanced (dashed-dotted line) and emerging economies in isolation (dashed line). The total number of observations in our sample is 1497, of which 783 are for advanced economies. In each case, the mass of observations is very much concentrated on the left. For the full sample, for instance, about 50 percent of the observations for the spread are below 1 percentage point. Still, there are considerable differences across the two country groups: more than 99.8 percent of observations are below 10 percentage points in the sample of advanced economies. The corresponding number is only 93.3 percent in the sample of emerging market economies. An alternative and widely considered indicator of debt sustainability are credit default swap (CDS) spreads.¹⁴ CDS are insurance contracts that cover the repayment risk of an underlying bond. The CDS spread indicates the annual insurance premium to be paid by the buyer. Accordingly, a higher perceived default probability on the underlying bond implies, *ceteris paribus*, a higher CDS spread. While well-suited to capture market assessment of debt sustainability, CDS data are generally only available after 2003 when a liquid market developed (see Mengle, 2007). To check the quality of our constructed spread measure, we compare it to yields of 5year CDS spreads. We find a correlation of 0.95 (see also Figure A.1).

Finally, in the last two columns of Table 2 we report the correlation of sovereign yield spreads with output growth and the growth of government consumption, respectively. It turns out that spreads are countercyclical in all countries, although sometimes the correlation is negligible. Instead, the correlation of spreads and government consumption growth varies across countries. It is negative in about two thirds of the countries, but is generally weak. Eventually, we seek to establish the co-movement of spreads and government consumption conditional on an exogenous variation in government consumption. In order to do so, we rely

¹²Denmark, Ireland, Greece, Spain, France, Italy, Portugal, Slovenia, United Kingdom, Belgium, Netherlands, Finland, and Sweden

¹³Croatia, Hungary, Poland, Turkey, Argentina, Colombia, South Africa, Ecuador, Chile, El Salvador, Malaysia, Thailand, Uruguay, and Peru.

¹⁴For instance, in a recent study, Longstaff, Pan, Pedersen, and Singleton (2011) find an important role of global factors in accounting for CDS spread dynamics of individual countries.

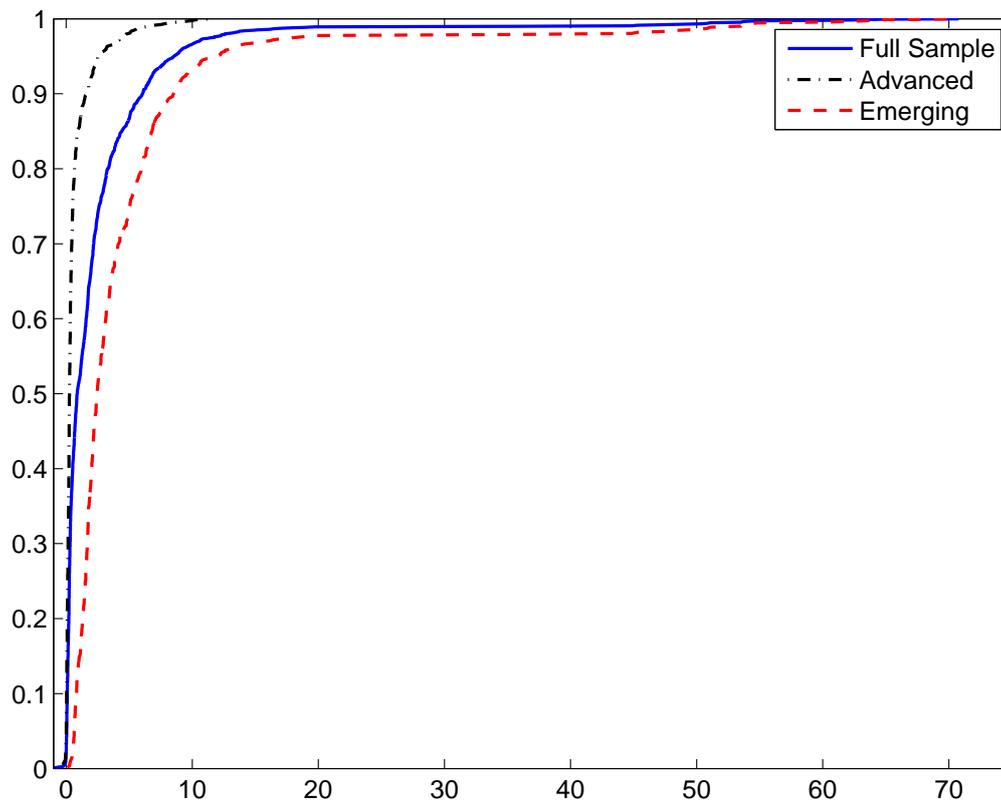


Figure 3: Sovereign yield spreads: empirical distribution function (CDF). Notes: horizontal axis measures spreads in percentage points. Vertical axis measures fraction of observations for which spread exceeds value on the horizontal axis. Solid line displays CDF for full sample (total number of observations: 1497), dashed-dotted line: advanced economies only (783), dashed line: emerging economies only (714).

on specific identification assumptions which are imposed within a particular econometric framework.

3 Econometric framework

In this section we discuss the econometric framework used to establish the effects of austerity on sovereign yield spreads. We first discuss identification and then turn to how we condition on the economic environment.

3.1 Identification

In terms of fiscal policy measures we focus on the dynamic effects of exhaustive government consumption for reasons of data availability. We obtain identification by assuming that, within a given quarter, government consumption is predetermined relative to the other variables included in our regressions. This assumption goes back to Blanchard and Perotti (2002) and has been widely applied in the empirical literature on the fiscal transmission mechanism. It is plausible, because government consumption is unlikely a) to respond automatically to the cycle and b) to be adjusted instantaneously in a discretionary manner by policy makers. To see this, recall that government consumption, unlike transfers, is not composed of cyclical items and that discretionary government spending is subject to decision lags that prevent policymakers from responding to contemporaneous developments in the economy.¹⁵

Still, influential work by Ramey (2011b) and Leeper, Walker, and Yang (2013), has made clear that identification based on the assumption that government spending is predetermined may fail to uncover the true response of the economy to a government spending shock whenever such shocks have been anticipated by market participants. Of course, the notion that fiscal policy measures are anticipated, because they are the result of a legislative process and/or subject to implementation lags is plausible.¹⁶ To address this issue, we follow Ramey (2011b) and Auerbach and Gorodnichenko (2013a), and consider a specification of our model where we include forecast errors of government consumption, rather than government spending itself. Given data availability, we show—for a subset of our sample—that results do not change qualitatively relative to our baseline case.

A number of promising alternatives to identify shocks to government spending have been pursued in the literature. Ramey (2011b) relies both on war dates and forecast errors for government spending. The latter are available for a subset of countries in our sample. Below

¹⁵Anecdotal evidence suggests that this holds true also in times of fiscal stress. For instance, in November 2009, European Commission (2009) states regarding Greece: “in its recommendations of 27 April 2009 ... the Council [of the European Union] did not consider the measures already announced by that time, to be sufficient to achieve the 2009 deficit target and recommended to the Greek authorities to “strengthen the fiscal adjustment in 2009 through permanent measures, mainly on the expenditure side”. In response to these recommendations the Greek government announced, on 25 June 2009, an additional set of fiscal measures to be implemented in 2009 ... However, these measures ... have not been implemented by the Greek authorities so far.” In fact, it appears that significant measures were put in place not before 2010Q1, see Greece Ministry of Finance (2010).

¹⁶Still, whether or not this invalidates the identification assumption is a quantitative matter (Sims, 2012). Results by Beetsma and Giuliodori (2011), Corsetti, Meier, and Müller (2012b), and Born, Juessen, and Müller (2013), for instance, suggest that the issue is of limited quantitative relevance as far as shocks to government spending are concerned.

we show that our key results are robust, once we adopt an identification scheme based on forecast errors and apply it to this subset of country observations. Another popular approach to identify fiscal shocks on a narrative basis. Following the work of Romer and Romer (2010) for the US, Devries, Guajardo, Leigh, and Pescatori (2011) have constructed a data set of fiscal shock for a large sample of OECD countries. However, these fiscal shocks are identified on a narrative basis with a view to being orthogonal to the business cycle. A large share of these shocks, however, reflects fiscal measures taken in order to reign in public debt or budget deficits. To the extent that sovereign yield spreads comove systematically with the latter, these “shocks” are not suited to investigate the effect of fiscal policy on sovereign yield spreads.¹⁷

Yet an alternative strand of the literature, following the lead of Alesina and Perotti (1995), identifies “fiscal adjustments” as episodes during which the cyclically adjusted primary deficit falls relative to GDP by a certain amount (see e.g. Alesina and Ardagna, 2010; IMF, 2010). In these studies, an episode of fiscal consolidation is assumed to take place over several years. Our focus, instead, is on the short run dynamics of spreads to cuts in government consumption. Finally, note that due to non-availability of appropriate tax data, we do not attempt to identify the effects of tax shocks.

3.2 Model specification

Our results are based on two alternative model specifications. Traditionally, the Blanchard-Perotti identification has been employed within a VAR context. More recently, it has also been used in panel VAR models, see, e.g., Ilzetzki, Mendoza, and Végh (2013) and Born, Juessen, and Müller (2013). Below we will also report estimates based on such a model.

In this case, the vector of endogenous variables includes three variables: the log of real government consumption, $g_{i,t}$, the log of real GDP, $y_{i,t}$, and sovereign yield spreads, $s_{i,t}$, measured in percentage points. In what follows, i denotes the country and t the time period. The VAR model is given by

$$\begin{bmatrix} g_{i,t} \\ y_{i,t} \\ s_{i,t} \end{bmatrix} = \boldsymbol{\mu}_i + \boldsymbol{\alpha}_i t + \sum_{k=1}^K \mathbf{A}_k \begin{bmatrix} g_{i,t-k} \\ y_{i,t-k} \\ s_{i,t-k} \end{bmatrix} + \boldsymbol{\nu}_{i,t}, \quad (3.1)$$

¹⁷Another strand of the literature employs sign restrictions to identify fiscal shocks, see Mountford and Uhlig (2009). This is not feasible in the context of our analysis, as there is no consensus on the sign of the responses to a fiscal shock as far as the variables in our sample are concerned.

where $\boldsymbol{\mu}_i$ and $\boldsymbol{\alpha}_i$ are vectors containing country-specific constants and time trends. The matrices \mathbf{A}_k capture the effect of past realizations on the current vector of endogenous variables. $\boldsymbol{\nu}_{i,t}$ is a vector of reduced form residuals. We estimate model (3.1) by OLS. Identification is based on mapping the reduced-form innovations $\boldsymbol{\nu}_{i,t}$ into structural shocks:

$$\boldsymbol{\varepsilon}_{i,t} = \mathbf{B}\boldsymbol{\nu}_{i,t}, \text{ with } \boldsymbol{\varepsilon}_{i,t} \stackrel{iid}{\sim} (0, \mathbf{I}).$$

In the present context, identifying shocks to government consumption under the assumption that it is predetermined boils down to equating the first element in $\boldsymbol{\nu}_{i,t}$ with a structural fiscal shock.¹⁸

Recently, local projections have been a popular tool to complement VAR analysis. As argued by Jordá (2005), local projections are more robust to model misspecification as they do not impose cross-equation restrictions as in sVAR models. Moreover, local projections prove highly flexible in accommodating a panel structure. Finally, and most importantly, they offer a very convenient way to account for state dependence—the focus of our analysis below. Earlier work by Auerbach and Gorodnichenko (2013a) has illustrated this in the context of fiscal policy. More specifically, they suggest a panel smooth transition autoregressive (STAR) model on which we rely below.

Defining the vector $X_{i,t} = [g_{i,t} \ y_{i,t} \ s_{i,t}]'$, the response of a variable $x_{i,t+h}$ at horizon h to $g_{i,t}$ can be obtained by locally projecting $x_{i,t+h}$ on time t government spending and a set of control variables/regressors. That is, the following relation is estimated:

$$\begin{aligned} x_{i,t+h} = & \alpha_{i,h} + \beta_{i,h}t + \eta_{t,h} \\ & + F(z_{i,t})\psi_{A,h}g_{i,t} + [1 - F(z_{i,t})]\psi_{B,h}g_{i,t} \\ & + F(z_{i,t})\Pi_{A,h}(L)X_{i,t-1} + [1 - F(z_{i,t})]\Pi_{B,h}(L)X_{i,t-1} + u_{i,t}. \end{aligned} \quad (3.2)$$

Here $\alpha_{i,h}$ and $\beta_{i,h}t$ are a country-specific constant and a country-specific trend, respectively. $\eta_{t,h}$ in turn captures time fixed effects, which we do not allow for in the VAR model above. $u_{i,t}$ is an error term with strictly positive variance. L denotes the lag operator. At each horizon, the response of the dependent variable to government spending is allowed to differ across regimes “A” and “B“, with the ψ -coefficients on the $g_{i,t}$ terms indexed accordingly. Similarly, $\Pi_{*,h}(L)$ is a lag polynomial of coefficient matrices capturing the impact of control variables in each regime. We estimate (3.2) using OLS, assuming that government spending

¹⁸As a practical matter, we impose a lower-triangular structure on \mathbf{B} , attaching no structural interpretation to the other elements in $\boldsymbol{\varepsilon}_{i,t}$.

is predetermined (see also Auerbach and Gorodnichenko, 2013b).

Conceptually it is convenient to distinguish two polar regimes which give rise to possibly different dynamics after a fiscal impulse. These polar cases are characterized by $F(z_{i,t})$ being equal to zero and one, respectively. It is quite unlikely, however, that actual economies operate in either of these regimes. Rather, they tend to be more or less close to one of the two. This is captured in the estimation, as the projection of the dependent variable at each horizon is a weighted average, whereby the fiscal shock as well as the control regressors are allowed to impact the dependent variable differently. The weights, in turn, are a function $F(\cdot)$ of an indicator variable $z_{i,t}$ which provides information of how close the economy is to one of the two regimes. By using this weighted average, all observations between the two polar help in identifying the two regimes.

In our estimation below we use lagged yield spreads $z_{i,t} = s_{i,t-1}$ as an indicator variable in order to measure how closely an economy operates to a regime of “fiscal stress”. Using the lagged value of the spread assures that the indicator is orthogonal to our identified government spending shocks. We weigh regressors on the basis of the country-group specific empirical CDF (see Figure 3 above). Formally, we have

$$F(z_{i,t}) = \frac{1}{N} \sum_{j=1}^N \mathbb{1}_{z_j < z_{i,t}} \quad (3.3)$$

where $\mathbb{1}$ denotes an indicator function and j indexes all country-time observations in the respective country group. As an alternative to the empirical CDF, one may assume a specific parametric function in order to map the indicator variable into specific weights.¹⁹ Using the empirical CDF, however, has two advantages. First, there are no degrees of freedom in specifying the transition function. Second, the polar cases are now given by states of the world that were actually obtained in-sample.

4 Results

We now turn to our estimates of the effects of government consumption cuts. Our main focus is the dynamic response of sovereign spreads to such cuts. Still, as argued above, because the adjustment of output is likely to be a key determinant of spreads, we also

¹⁹Auerbach and Gorodnichenko (2012) use a logistic cumulative density function $F(z_{i,t}) = \frac{\exp(-\gamma z_{i,t})}{1 + \exp(-\gamma z_{i,t})}$ as their transition function so that $Prob(z < \bar{z}) = F(\bar{z})$. The free parameter γ was chosen using extraneous evidence that the US is in recession 20% of the time.

report the response of output alongside that of government consumption and the spread. We normalize the size of the initial shock such that the cut in government consumption is equal to one percent of GDP.

In a first subsection, we report results for a linear model, not accounting for possible state dependence. Instead we contrast results for local projections with those obtained from a VAR model. Afterwards we document that results differ considerably, depending on the state of the economy. Recall that, in each instant, we obtain identification by assuming that government spending is predetermined within a given quarter.

4.1 Linear model

Figure 4 shows results for the baseline case: the specification which does not allow for state dependent effects, but constrains the effects of fiscal shocks to be linear. The panels in the upper row show impulse responses based on estimates obtained from local projections (LP). Here and in the following the horizontal axis measures time in quarters, while the vertical axis measures the deviation from the pre-shock path. For real quantities it is measured in percent of trend output, for the spread it is measured in basis points. Solid lines represent the point estimates, while shaded areas indicate 90% confidence bounds. The second row shows the impulse responses obtained from the estimated panel vector autoregression (VAR) model.²⁰ Note that the horizon for which we report impulse responses differs for the LP and the VAR model, as extending the horizon in the former case comes at the expense of degrees of freedom. The VAR results for the first 6 quarters are very similar to the LP results, both from a qualitative and a quantitative point of view.

The first column of Figure 4 shows the dynamic adjustment of government consumption over time. After the initial 1% of GDP cut, government consumption remains depressed for an extended period, but eventually returns to its pre-shock level, as evidenced by the VAR results shown in the second row. The response of GDP is displayed in the panels of the second column. It declines by about 0.3 percentage points on impact, declines further and reaches a trough response of about -.6 percent of GDP after about 3 quarters. Given that we normalize the initial cut in government consumption to -1 percent of GDP, these estimates can be interpreted as estimates of the government spending multiplier on output (impact and peak-to-impact, respectively). Our estimates fall in the range of values frequently reported in the literature, if perhaps somewhat at the lower end (see e.g. Ramey, 2011a).

²⁰LP confidence bounds are based on Driscoll and Kraay (1998) cross-sectional correlation robust standard errors, VAR confidence bounds are bootstrapped using 100 bootstrap samples.

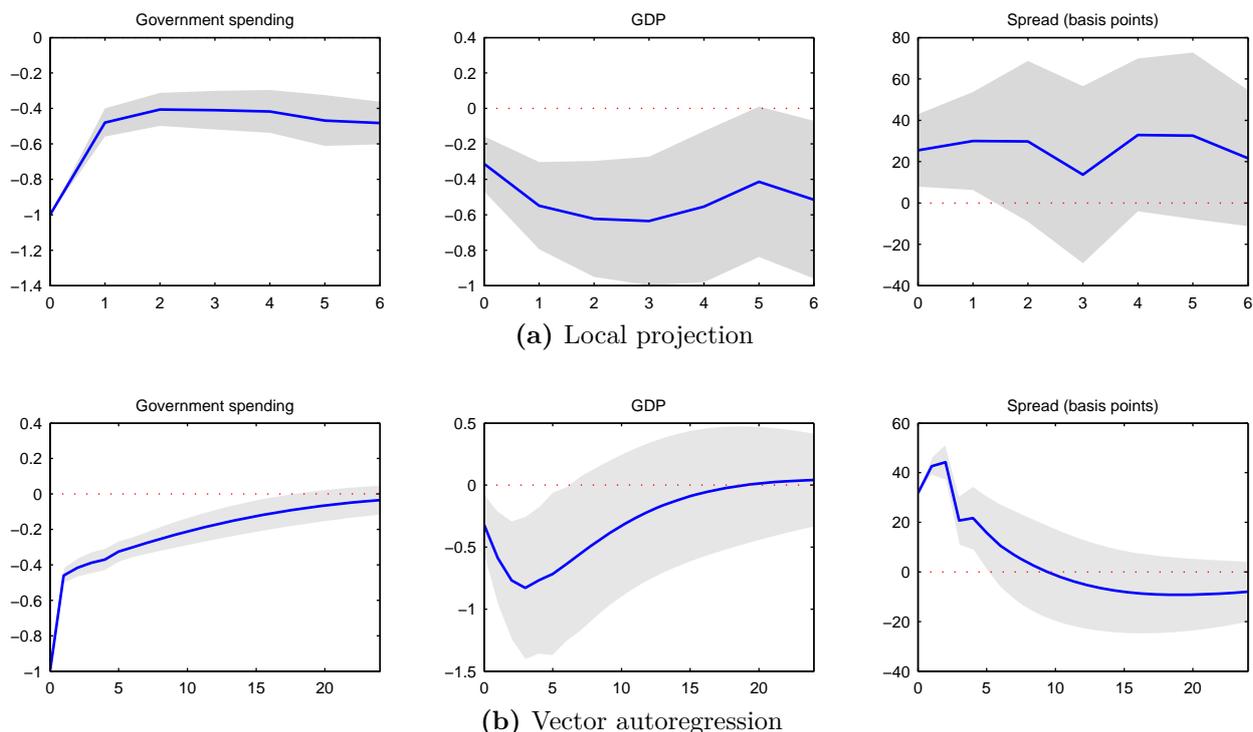


Figure 4: Dynamic response to a government spending shock in linear model. Notes: confidence bounds represent 90 percent standard errors. Horizontal axis represent quarters. Vertical axes represent deviation from pre shock level in terms of trend output and basis points (spread). Top panels show results based on local projections, bottom panels show results based on VAR.

Finally, in the third column, we present estimates for the dynamic response of spreads to the cut in government consumption. Here we find that spreads do, in fact, increase in response to the spending cut. The impact response varies between 25 basis points (LP) and 30 basis points (VAR). For both specifications we find a maximum effect of about 40 basis points. The VAR response shows that the spread returns to its pre-shock level, after mildly undershooting it for an extended periods. It thus appears that austerity doesn't pay off: spending cuts fail to reassure investors about the sustainability of public finances.²¹

4.2 Fiscal stress vs benign times

The above result obtains for the entire sample, possibly masking heterogeneity of economic circumstances which may matter for how spreads respond to austerity, both across

²¹It is also worth pointing out that the continuous movements of spreads over time do not violate market efficiency. Spreads move with fundamentals in basic models of sovereign default (see, e.g., Arellano, 2008).

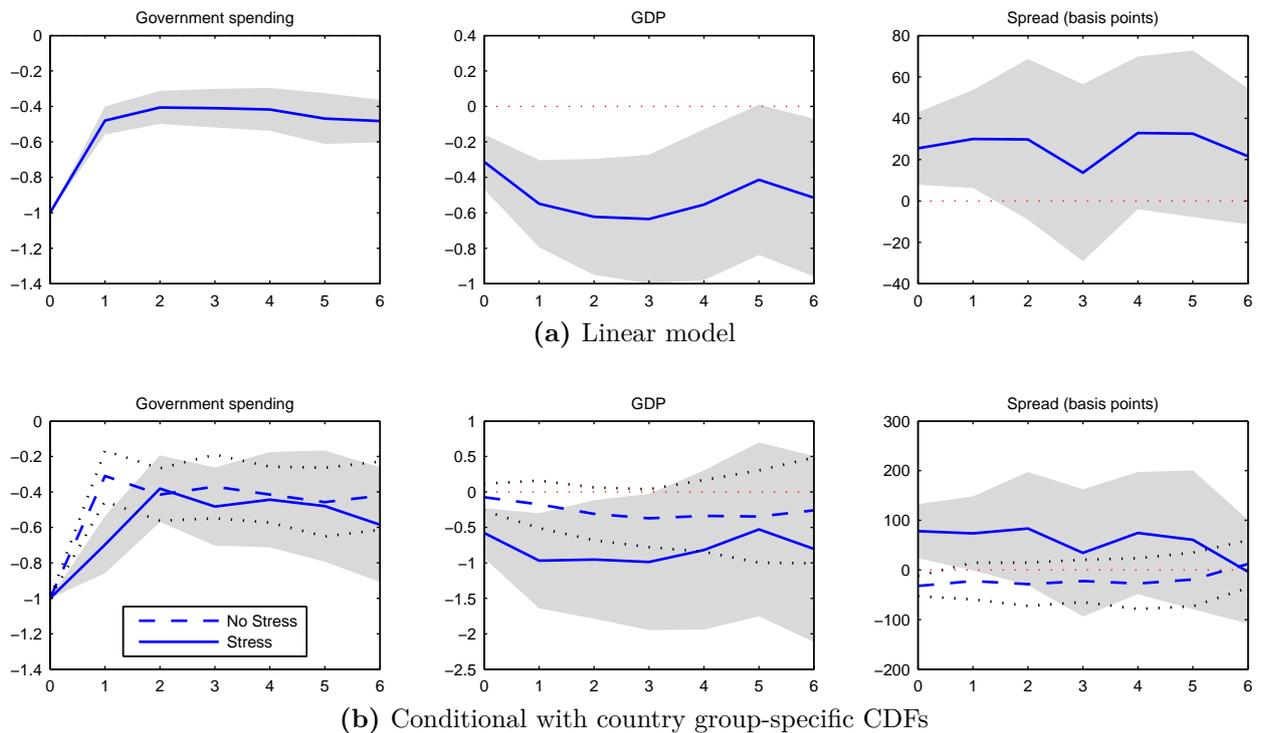


Figure 5: Dynamic response to a government spending shock derived from local projections (two lags). All standard errors are clustered robust 90% standard errors.

time and countries. In particular, earlier studies show that fiscal policy may affect the economy differently in “bad times” (Bertola and Drazen, 1993; Perotti, 1999). And indeed, recent evidence established by Corsetti, Meier, and Müller (2012a), Auerbach and Gorodnichenko (2013a) and Ilzetzi, Mendoza, and Végh (2013) suggests that the government spending multiplier on output tends to be relatively low, if debt is high. This is particularly relevant, as austerity is often enacted in response to concerns about the sustainability of debt. However, as discussed above, public debt *per se* is an insufficient statistic to assess the sustainability of public finances, because fiscal capacity varies strongly with a number of country-specific factors. Instead, sovereign yield spreads provide more comprehensive information regarding the extent of “fiscal stress”, both because of the underlying—arguably broader—assessment of financial market participants and, not least, because of the immediate budgetary consequences.

In what follows we therefore estimate the non-linear model (3.2) relying on the spreads as an indicator variable and their empirical CDF (3.3) as weighting function. We thus allow the effects of spending cuts to differ depending on whether they are in a regime of fiscal stress, evidenced by high spread or not (“benign times”). Recall, however, from

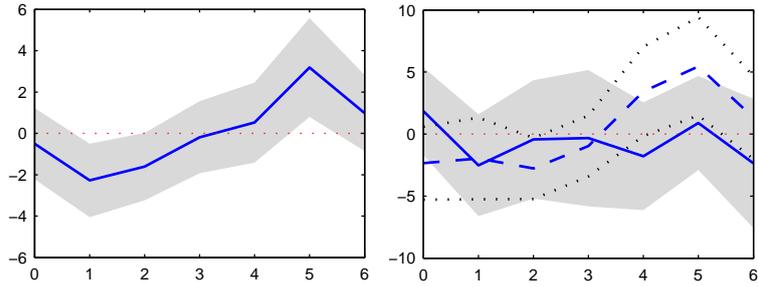


Figure 6: Confidence: dynamic response to a government spending shock. Left panel: linear model (LP). Right panel: responses conditional of fiscal stress (solid lines) and benign times (dashed lines). Confidence, based Ifo World Economic Survey, pertains to expectations regarding economic conditions in the next 6 months.

our discussion in section 3 that we allow for a smooth transition across regimes, as we weigh observations according to the relative size of the spread during a given country-time observation. For the baseline specification, in order to account for fundamental heterogeneity across the set of advanced and emerging economies, we use the empirical CDF obtained for each country group in isolation.

The second row of Figure 5 reports the results for the baseline specification, contrasting it to the results for the linear model (reproduced in the first row of Figure 5). Solid lines represent point estimates for the regime of fiscal stress, with shaded areas indicating 90 percent confidence bounds. Dashed lines represent the estimates for the other polar case, the response to austerity during benign times. Results are rather stark: the dynamic adjustment of the economy under fiscal stress resembles those obtained for the unconditional estimates very closely. Hence, fiscal stress episodes apparently dominate the overall sample. Only from a quantitative point of view, we find that the effects under fiscal stress differ from the baseline case: they are considerably magnified. The point estimate for the multiplier now reaches a value of about 0.7, while spreads rise to up to approximately 80 basis points in response to a cut of government consumption.

The effects of austerity in benign times, on the other hand, differ considerably from those obtained from the unconditional estimates. We now find the output effect not significantly different from zero. Importantly, our estimates also suggests that spreads decline in response to cuts in government consumption, provided that the economy enjoys more benign times. In this case spreads come down by about 50 basis points.

The adjustment of the economy to the fiscal shock thus differs considerably, depending on the two regimes. Earlier research on the consequences of fiscal consolidations has argued that its impact on “confidence” is crucial (see, for instance, the discussion in Perotti, 2013).

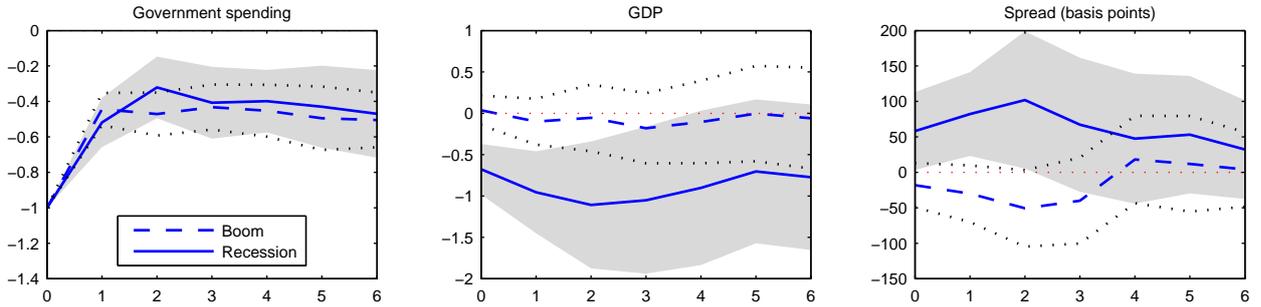


Figure 7: Dynamic response to a government spending shock during booms and recessions. Notes: output gap used as indicator variable.

Bachmann and Sims (2012) find that confidence responds strongly to fiscal shocks during periods of economic slack. To gain a better understanding of what mechanism may drive our results, we therefore estimate the dynamic response of confidence to a fiscal shock. For this purpose we rely on the Ifo World Economic Survey (WES), which surveys a number of experts for all countries in our sample.²²

Figure 6 displays the results. In the left panel we show results obtained for the linear model. As in Bachmann and Sims (2012) we find that confidence is fairly flat after a fiscal shock, although there is a marginally significant increase after about 5 quarters. Results do differ, however, once we condition on fiscal stress (right panel). In times of fiscal stress (solid lines) confidence is still unresponsive with respect to the spending cut. In benign times, in contrast, confidence tends to improve markedly after about a year after a spending cut. These findings are consistent with the notion that austerity is less harmful to economic activity whenever it is associated with an improvement of confidence. In our setup this coincides also with a decline in yield spreads.

Times of fiscal stress are mostly likely times of low output growth, for reasons discussed above. Of course, the converse does not necessarily hold: a recession may not give rise to fiscal stress if public finances are in good shape. Still, to put our results into perspective, it is useful to assess to what extent the effects of austerity on spreads change with the state of the business cycle. For this purpose we compute, following Auerbach and Gorodnichenko (2013a), a measure of the output gap.²³ We use it as an indicator variable and compute the empirical CDF as in the case of sovereign yield spreads (see Figure A.2 in the appendix).

²²Respondents are asked to classify their expectations for the next six months using a grid ranging from 1 (deterioration) to 9 (improvement). 5 indicates that expectations are “satisfactory”.

²³First, we compute a five-quarter moving average of the first difference of log output. The resulting series is then filtered using an Hodrick-Prescott filter with smoothing parameter $\lambda = 10,000$.

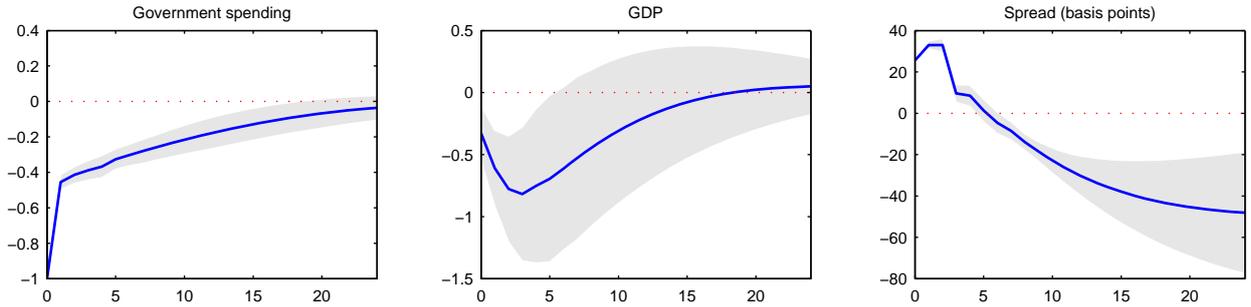


Figure 8: Responses to cut in government consumption with possibly long-run effect on spread. Notes: estimates based on VAR model, with spread entering in first differences.

Figure 7 shows the results. We obtain a pattern of responses quite comparable to the one obtained once we condition on fiscal stress. During recessions, as in times of fiscal stress, the multiplier is relatively large and spreads rise strongly in response to austerity. Conversely, during booms, as in fiscally benign times, the multiplier is basically zero and spreads decline in response to austerity. Perhaps surprisingly, while conditioning on fiscal stress and recessions yields very similar results, we find that the overlap of stress and recession episodes is far from complete. In particular, the correlation of the empirical CDF which are used in the projection as weights is only moderate (see Table A.1 in the appendix).

A robust finding of our analysis is that spending cuts during times of fiscal stress do not pay off: they induce spreads to rise. The same result obtains in the linear model. However, our focus so far has been on the short-run, that is, the first six quarters after the shock. Results for the VAR model, reported in the bottom row of Figure 4, instead, suggest that spreads may decline in the medium-term. In the long-run spreads return to the pre shock level. This is unsurprising, as the spread enters the VAR model in levels. We therefore consider an alternative specification in which the spread enters the VAR in first differences. Figure 8 displays the impulse responses. The right panel is of particular interest. It shows the cumulative response of yield spreads. The short run dynamics are comparable to our baseline specification, but the medium-term dynamics are quite different. Spreads fall considerably. In the long-run, we find a permanent decline of about 50 percentage points.

4.3 Sensitivity analysis

We explore the robustness of our findings across a range of alternative specifications and sample periods. A first set of experiments is aimed at exploring issues pertaining to

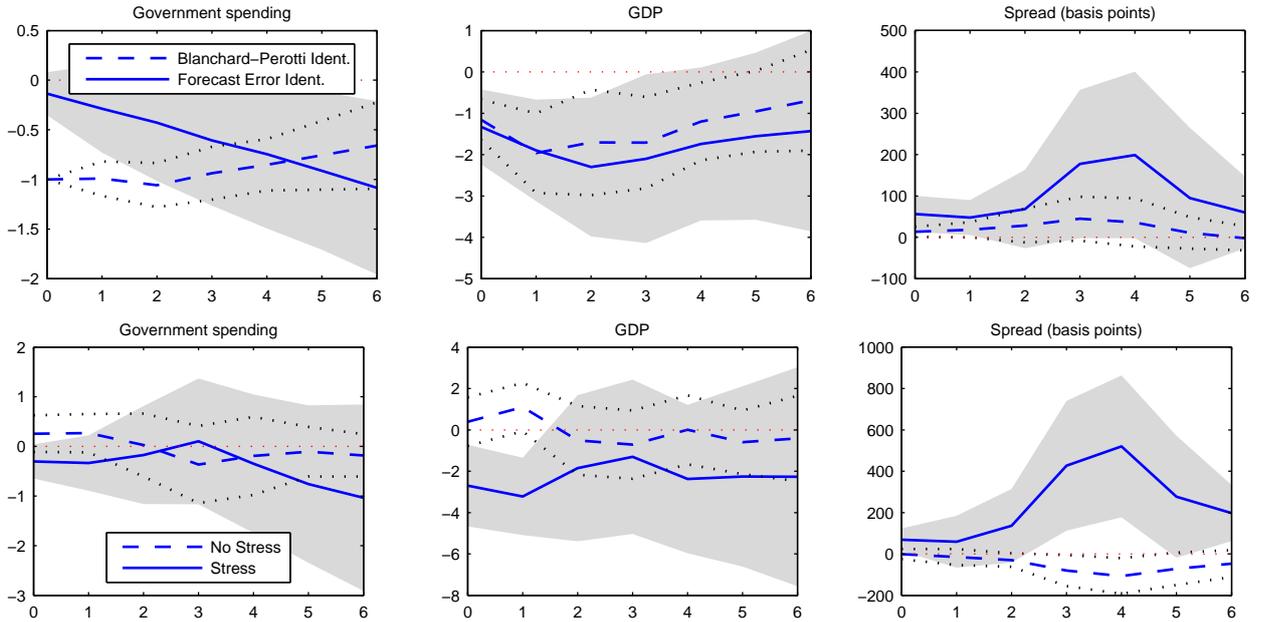


Figure 9: Dynamic response to a government spending shock derived from local projections (two lags). Solid line: Identification using OECD forecast errors. Dashed line: Blanchard/Perotti identification. All standard errors are clustered robust 90% standard errors. The country sample is restricted to observations with available OECD forecast errors.

identification. Importantly, as discussed above, under the Blanchard-Perotti identification approach, news and realizations of fiscal shocks necessarily coincide. To the extent that fiscal shocks become known prior to an actual change in fiscal variables, estimates may be biased (Ramey, 2011b; Leeper, Walker, and Yang, 2013). To gauge the impact of possibly anticipated government spending shocks on our results, we turn to the OECD Economic Outlook data set, which contains semiannual observations for the period from 1986 to 2013 for an unbalanced panel of OECD countries. It contains explicit forecasts for government consumption spending, prepared by the OECD in June and December of each year, that is, at the end of an observation period.²⁴

Including the forecast error for spending in the local projection model (3.2), rather than government spending allows us to better identify the effects of *unanticipated* spending shocks in the face of exogenous, but *anticipated* changes of government spending. Specifically, we

²⁴As discussed in detail by Auerbach and Gorodnichenko (2012), these forecasts have been shown to perform quite well. Auerbach and Gorodnichenko (2012) use these data to estimate government spending multipliers on the basis of local projections, contrasting results for recessions and booms. Beetsma and Giuliodori (2011) also control for anticipation effects when estimating the effects of government spending shocks. They consider annual data and include the budget forecast of the EU commission in their regression.

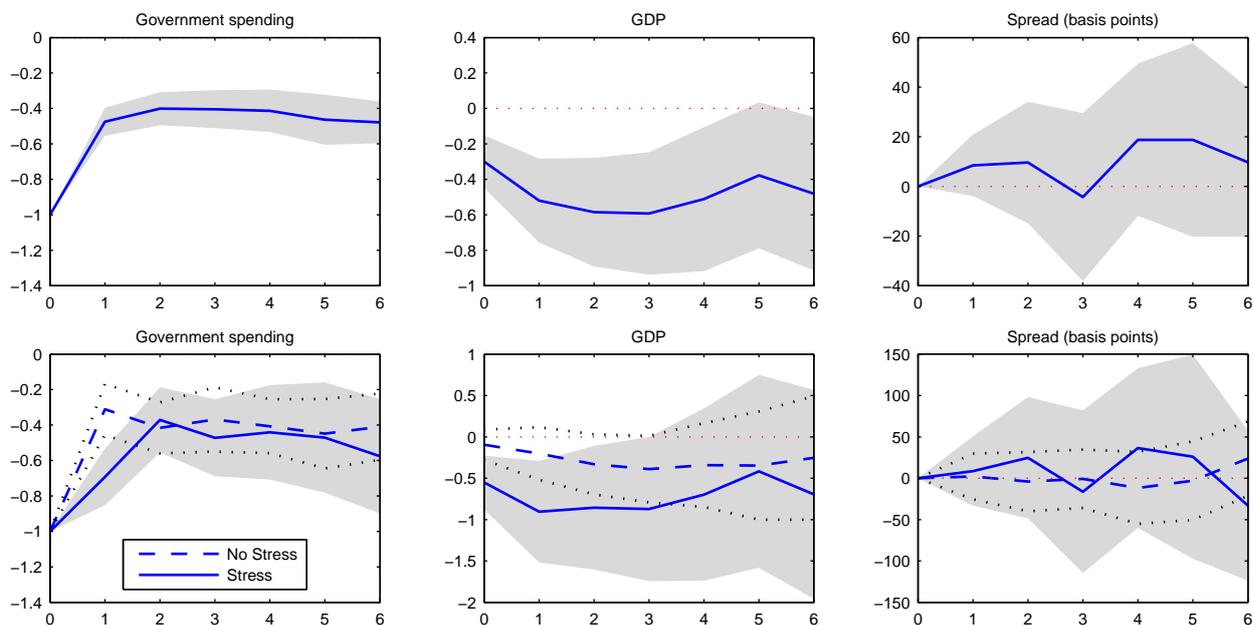


Figure 10: Dynamic response to a government spending shock derived from local projections (two lags). Contemporaneous spread included in control vector.

replace the level of government consumption with the period- t forecast error of the growth rate of government spending.²⁵ Figure 9 displays the results, obtained for the sample for which government spending forecasts are available. In the first row we contrast results based on forecast errors (solid lines) and those obtained under our baseline approach (dashed lines). It turns out that explicitly accounting for anticipation does not alter results very much (see also Beetsma and Giuliodori, 2011; Corsetti, Meier, and Müller, 2012b; Born, Juessen, and Müller, 2013). In the second row we show that the main result of our analysis, the differential effect of spending cuts during times of fiscal stress and benign times, also obtains, once identification is based on forecast errors.

Our results are based on the identification assumption that government spending is predetermined relative to output and yield spreads. This assumption is plausible to the extent that decision lags prevent an immediate discretionary policy response to either the cycle or fiscal stress. Still, in the light of our finding that spending cuts raise spreads, one may worry that we actually pick up reverse causation: spending falls as spreads rise—despite the fact that a within-quarter adjustment of spending is unlikely due to institutional constraints. We thus consider an alternative specification, where we include the contemporaneous value

²⁵We use growth rates rather than levels, because the base year used by the OECD changes several times during our sample period.

of the spread in the control vector. At the same time we impose the restriction that spreads do not contemporaneously respond to government spending shocks. Figure 10 displays the results, both for the linear model (top row) and the smooth transition model (bottom row). We find that results are qualitatively unchanged relative to those obtained for the baseline specification.

In a second set of experiments we explore the robustness of our results with respect to changes in the composition of the sample. First, we consider the full sample, but, in contrast to the baseline specification, use a common empirical CDF as a weighting function. Figure A.3 in the appendix shows the results which are quite similar to those obtained for the baseline specification. Next, we exclude the Great Recession from our sample, that is, we consider only observations up to the second quarter of 2007. Figure A.5 in the appendix shows the results based on the LP approach, distinguishing unconditional estimates from those obtained once we condition on fiscal stress and the business cycle. Contrasting the results with those for the full sample, we conclude that results are not driven by the Great Recession.

We also assess the robustness of our results regarding the role of fiscal stress through a number of sample splits, obtaining results for a sample which includes only euro area countries, for a sample of euro area periphery countries hit hardest by the crisis (Greece, Ireland, Italy, Portugal, Slovenia, Spain), and for a sample of the remaining euro area countries. Results, shown in Figure A.6, tend to be qualitatively similar to those obtained for the full sample—notably in terms of the differential impact of fiscal stress. The same holds for sub-samples comprising advanced and emerging economies only, see Figure A.7. As a caveat, however, we note that there are sizeable differences in some instances, reflecting perhaps also a strong decline in sample size.

5 Conclusion

Does austerity restore market confidence, bringing about a reduction in sovereign yield spreads? In pursuing this question, this paper makes two distinct contributions. First, we set up a new data set which contains data on sovereign yield spreads for 26 emerging and advanced economies. We assemble quarterly observations from 1990 to 2013, not only for spreads, but also for government consumption and output. A first look at the data allows us to establish a number of basic facts. First, while there is a large variation in yield spreads, both across time and countries, yield spreads are moderate for the largest

part of our sample. Second, yield spreads are strongly countercyclical. The correlation of yield spreads and current output growth is negative in almost all countries of our sample. Third, there is no systematic correlation pattern emerging for yield spreads and government consumption.

As a second contribution, we assess how yield spreads react to austerity measures. If we do not condition on the state of the economy, we find that a cut of government consumption raises sovereign yield spreads in the short run. At the same time output declines considerably. A cut of government consumption by one percent of GDP raises spreads by 40 to 50 basis points and reduces economic activity by about 0.6 percentage points. It turns out that these effects are driven by episodes of fiscal stress. If we condition estimates on fiscal stress, captured by high yield spreads, we find that spending cuts have an even stronger effect on spreads: they increase by about 80 to 90 basis points. Similarly, the adverse output effects are also amplified in this case.

Instead, if the economy enjoys more benign times, spending cuts pay off in that they bring about a sizeable reduction of yield spreads (about 50 basis points). In this case, output is hardly affected by the cut of government consumption. Moreover, for the linear model, we also find that spreads tend to decline in the long run once economic activity has fully recovered. In sum, the data reveal a very robust pattern: yield spreads tends to move negatively with output—both, unconditionally and conditional on fiscal shocks. Hence, to the extent that austerity impacts economic activity adversely, it likely fails to bring about a reduction in yield spreads. These findings are consistent with the view that financial markets are primarily concerned with output growth.

Austerity may pay off in the long-term or if it is implemented during benign times. Under adverse fiscal conditions, instead, it may be beneficial to delay austerity measures. In this case, in order to reassure markets about the sustainability of public finances, one may rather enact policies directed towards boosting economic activity. While taken at face value, our results suggests that even expansionary fiscal policies may be beneficial in this regard, we caution against such conclusions, because of the possibly adverse long-term implications. These, in turn, warrant further investigation.

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A Appendix

Table A.1: Descriptive statistics: indicators

Country	$\rho(F(z_{i,t}^{stress}), F(z_{i,t}^{ygap}))$	$\text{mean}(F(z_{i,t}^{stress}))$	$\text{mean}(F(z_{i,t}^{ygap}))$
Belgium	-0.32	0.51	0.50
Denmark	-0.23	0.59	0.43
Finland	-0.22	0.44	0.52
France	-0.12	0.35	0.52
Greece	-0.54	0.62	0.47
Hungary	-0.17	0.25	0.53
Ireland	-0.21	0.48	0.50
Italy	-0.03	0.60	0.49
Netherlands	-0.35	0.31	0.50
Poland	-0.39	0.31	0.48
Portugal	-0.17	0.48	0.48
Slovenia	-0.56	0.73	0.57
Spain	-0.28	0.46	0.50
Sweden	0.06	0.59	0.47
United Kingdom	-0.41	0.49	0.56
Argentina	-0.25	0.84	0.50
Chile	-0.74	0.25	0.50
Colombia	-0.33	0.54	0.49
Ecuador	-0.16	0.90	0.48
El Salvador	-0.59	0.57	0.54
Malaysia	-0.45	0.27	0.51
Peru	-0.38	0.57	0.48
South Africa	-0.54	0.40	0.51
Thailand	-0.36	0.24	0.51
Turkey	-0.28	0.65	0.51
Uruguay	0.12	0.60	0.48

Notes: $F(z_{i,t}^{stress})$ denotes the values of the country group-specific empirical CDF of the lagged spread; $F(z_{i,t}^{ygap})$ denotes the empirical CDF of the the smoothed output gap, computed as the z-scored deviation of the 5 quarter moving average of the output growth rate from its HP-filtered trend ($\lambda = 10,000$). First column: correlation between the two CDFs. Negative values indicate that fiscal stress is higher when the output gap is more negative. Second column: average value of the fiscal stress CDF for the respective country. Last column: average value of the output gap CDF for the respective country.

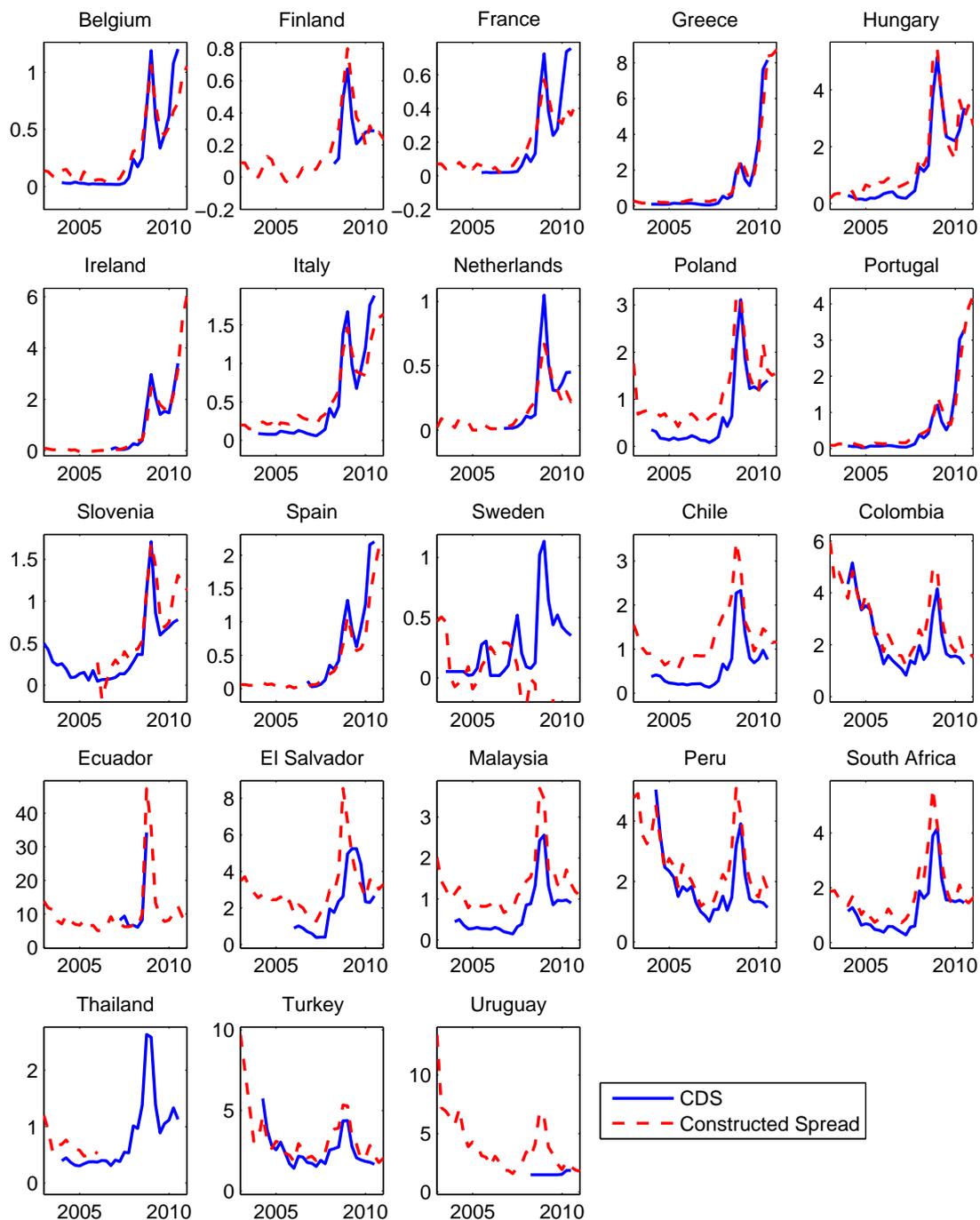


Figure A.1: Comparison CDS vs constructed spreads for selected countries. Blue solid line: 5year CDS yields on government bonds. Red dashed line: yield spread series on government bonds, constructed as described in the text.

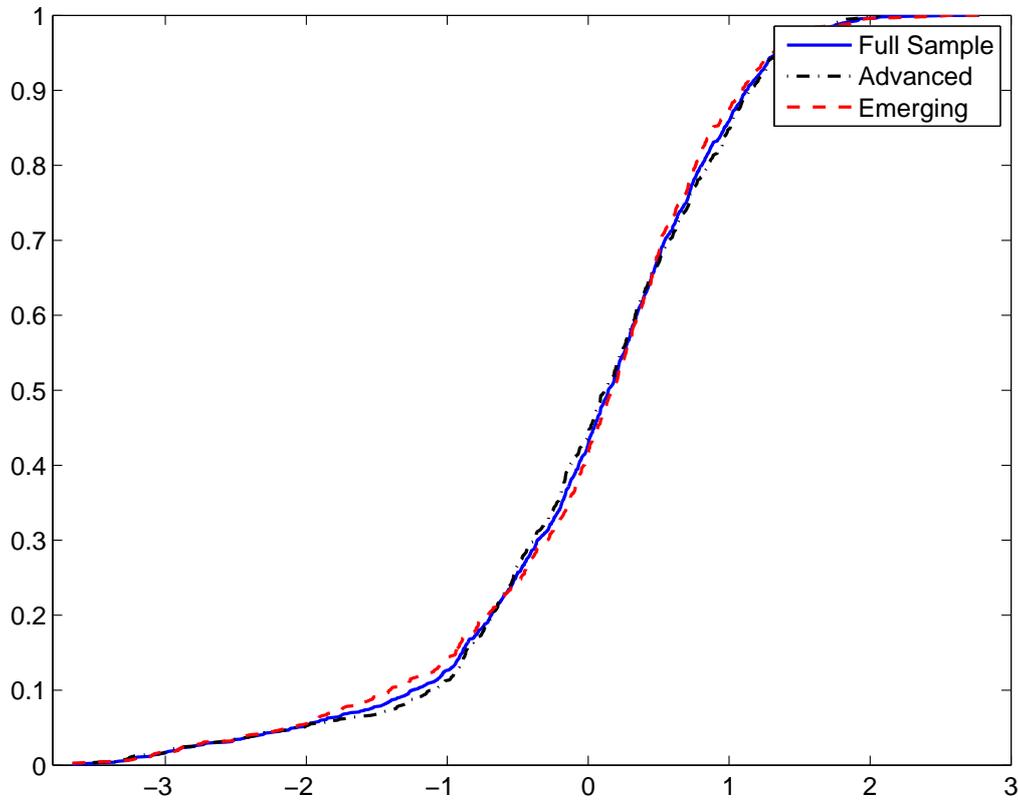


Figure A.2: Smoothed output gap: empirical distribution function (CDF). Notes: horizontal axis measures the smoothed output gap, computed as the z-scored deviation of the 5 quarter moving average of the output growth rate from its HP-filtered trend ($\lambda = 10,000$). Vertical axis measures fraction of observations for which spread exceeds value on the horizontal axis. Solid line displays CDF for full sample (total number of observations: 1497), dashed-dotted line: advanced economies only (783), dashed line: emerging economies only (714).

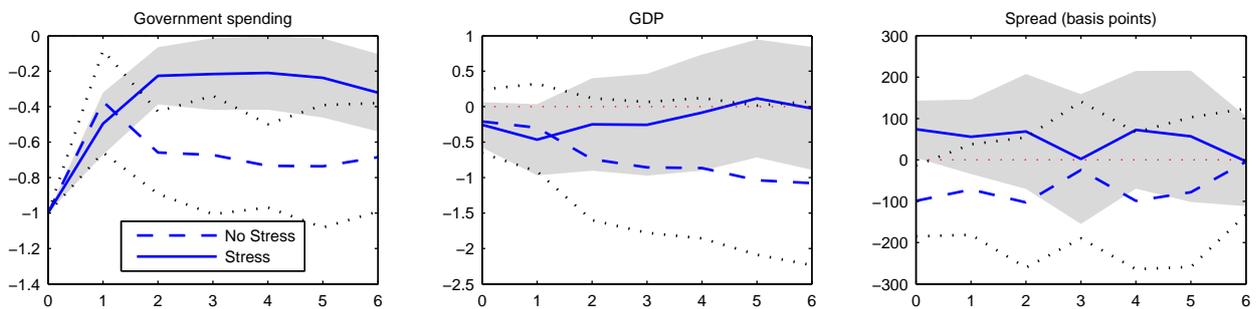


Figure A.3: Dynamic response to a government spending shock: common CDF for both country groups.

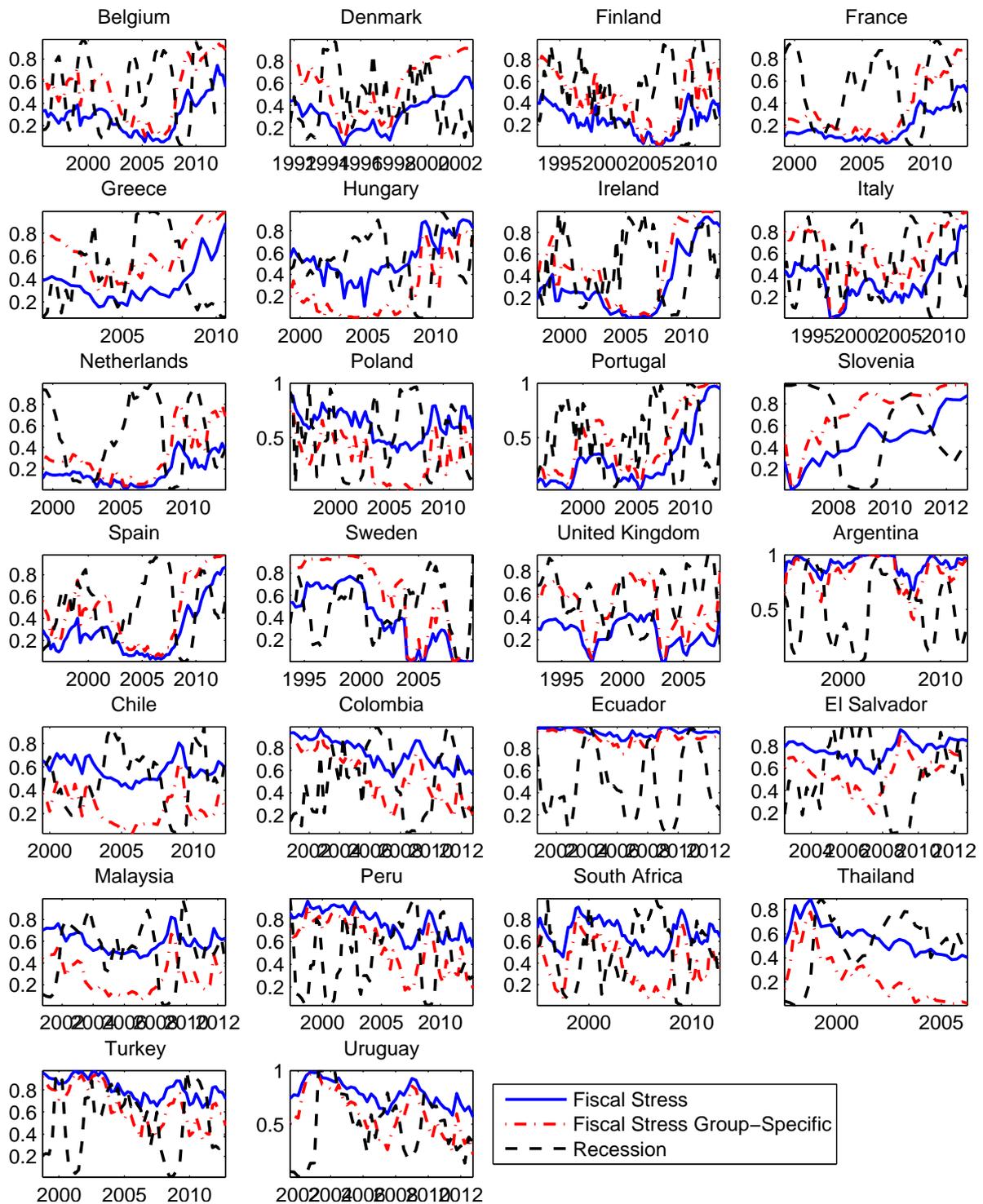


Figure A.4: Empirical CDF values for spreads and smoothed output gaps.

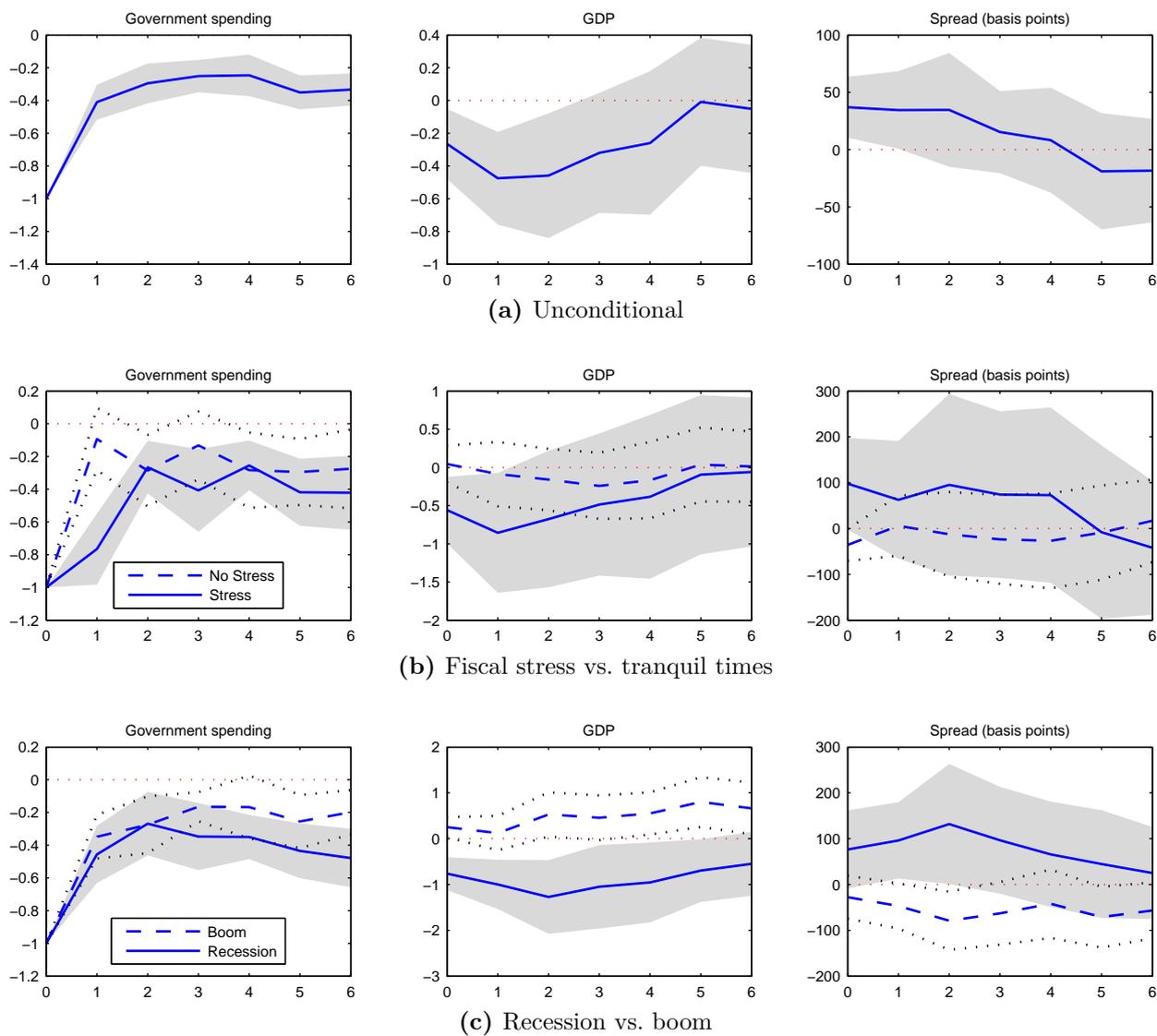


Figure A.5: Dynamic response to a government spending shock derived from local projections (two lags) when excluding the Great Recession. All standard errors are clustered robust 90% standard errors.

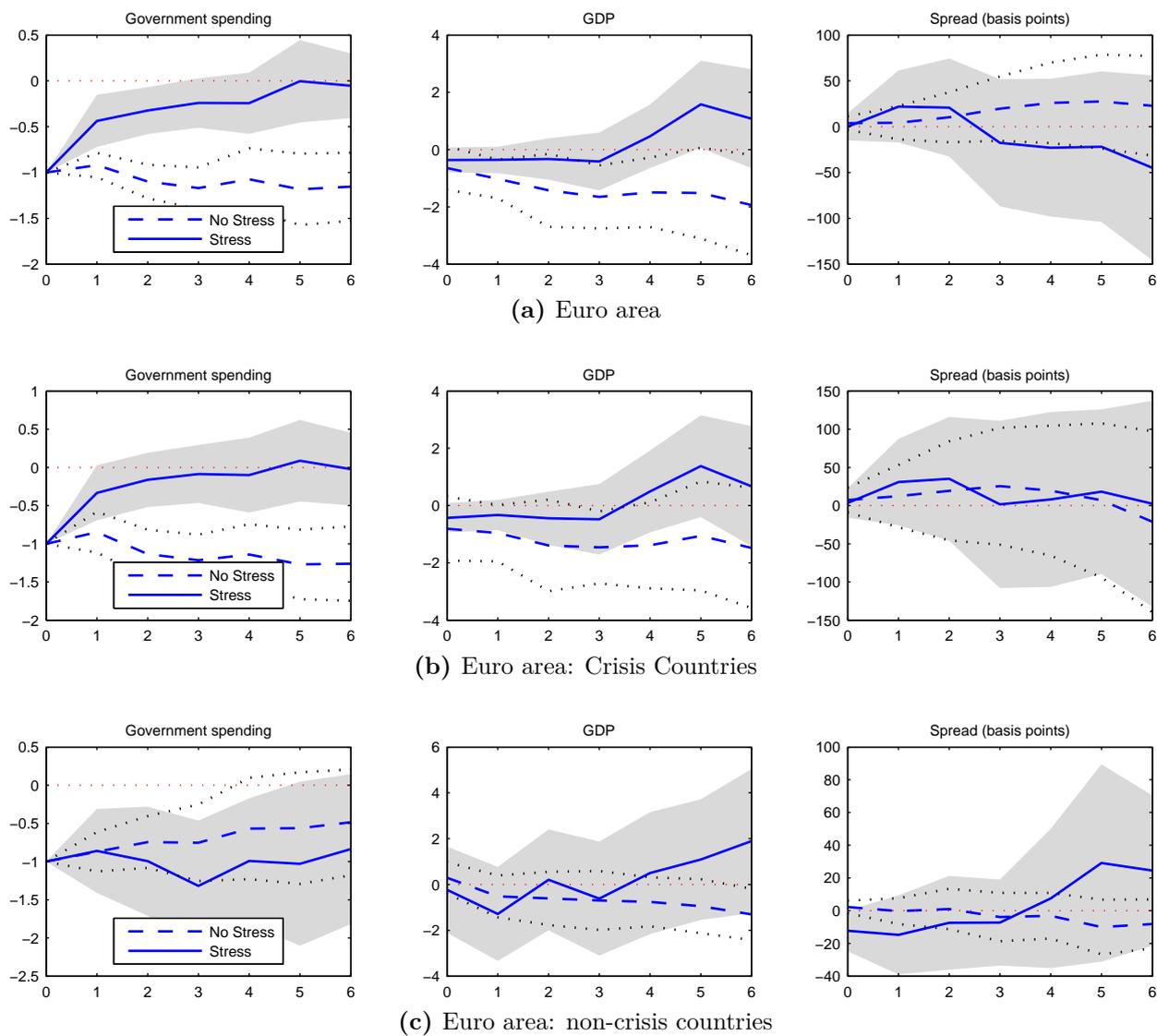


Figure A.6: Dynamic response to a government spending shock derived from local projections for the whole sample and different euro area countries. All standard errors are clustered robust 90% standard errors.

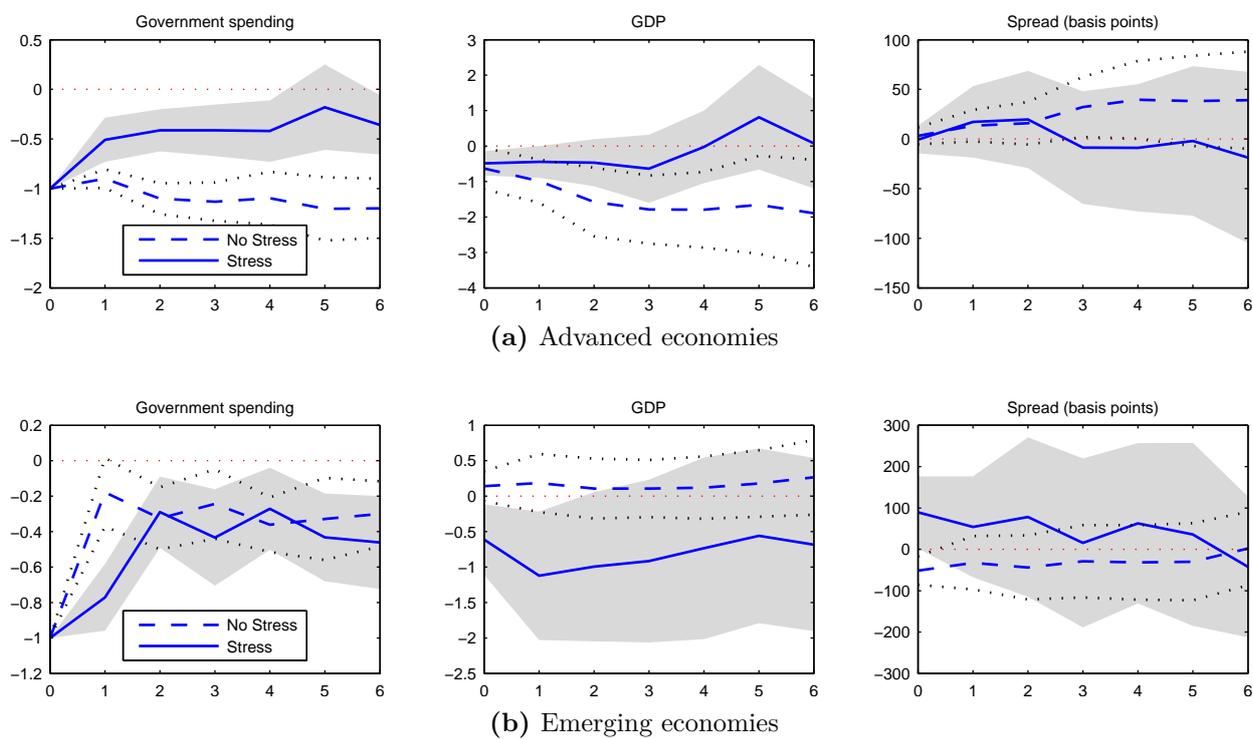


Figure A.7: Dynamic response to a government spending shock derived from local projections for emerging and developed countries. All standard errors are clustered robust 90% standard errors.