Essays on Government Spending and Employment

By

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Abstract

The essays that follow analyze recent developments in government spending and labor markets. In Chapter 1, I examine the "twin deficits" hypothesis in a new light by pooling data from highly developed open economies to create a representative global economy. Previous work on twin deficits has yielded mixed evidence supporting and refuting this theoretical explanation for the comovement of government budget and current account deficits. I estimate a vector autoregression (VAR) model with panel data using Bayesian methods to determine the effects of shocks to government spending on current account balances and real interest rates using various groups of nations and time periods. The results indicate that increases in government spending do not lead to increases in the current account deficit. My findings do not support the predictions of the twin deficits hypothesis.

Motivated by the results in Chapter 1, Chapter 2 investigates the effects of government spending shocks on output, consumption, employment, and interest rates. To address empirical challenges related to fiscal policy lags, data availability, and simultaneity, Chapter 2 uses the narrative approach to record announcements of changes in military, national security, and counterterrorism spending in the form of a defense and civil defense news variable called *defnews*. This variable captures variation in defense and civil defense spending since September 11, 2001, a period that provides a natural experiment setting in which frequent and meaningful changes in defense and civil defense spending occurred independently of global and national business cycles. I implement a unique and extensive data gathering process to combine data from six nations and use a panel VAR

model with Bayesian estimation methods to show that *defnews* correctly identifies changes in total government expenditures. The paper proposes a new set of stylized facts describing the effects of changes in government purchases that support the predictions of the neoclassical model and describe a value of the fiscal multiplier that is non-positive and close to zero.

Chapter 3 investigates changes in the relationship between unemployment duration and unemployment rates in the United States. Although unemployment rates have trended downward in the United States since the 1970s, mean unemployment duration has risen. Over this same period, the dynamics of female labor supply have changed dramatically: large numbers of women have entered the labor force, women's wages have risen, and wage elasticities have fallen. I use Current Population Survey data to match individuals to their spouses and look at how family labor supply (specifically, spouse wages and spousal employment) has affected unemployment duration. This paper is coauthored with Professor Donna Ginther from the University of Kansas and Melinda Pitts from the Federal Reserve Bank of Atlanta.

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I must also recognize the role my parents, Jane and Steve, and my brother, Chris, have played during my time in the Ph.D. program. Their advice, encouragement, and open ears helped me navigate the peaks and valleys of coursework and research. I dedicate this work and the time and effort that went into it to them.

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Chapter 1

Examining the Twin Deficits Hypothesis in a Macroeconomic Panel

1.1 Introduction and Relevance of the Twin Deficits Hypothesis

The twin deficits hypothesis has been the subject of persistent macroeconomic inquiry and debate over the last thirty years. Basic economic theory tells us that increases in government spending and inelastic demand for government bonds cause bond prices to fall when supply rises to finance fiscal expansion. According to Mundell (1963), if capital is perfectly mobile and capital flows are unrestricted, rising interest rates generate capital inflows that lead to domestic currency appreciation and deterioration of the trade balance. This is the traditional "twin deficits" story that is taught in introductory economics classes.

Between 1980 and 1986, as the U.S. federal budget deficit rose from \$59 billion to \$221 billion, the U.S. current account balance moved from a surplus of \$2.3 billion to a deficit of \$147.2 billion. Between 2001 and 2004, the U.S. federal budget deficit rose from -\$128 billion to \$413 billion while the U.S. current account deficit rose from \$395 billion to \$634 billion. In Germany (between 2003 and 2007) and Japan (between 1996 and 2000), increases in government spending were matched by simultaneous increases in current account surpluses. These data appear to be at odds with each other, and with basic economic theory.

In recent years the broader issue of government budget deficits drew international attention due to large fiscal stimulus measures enacted following the financial crisis (e.g., the American Reinvestment and Recovery Act and similar proposals across European Union member states). In the U.S., fiscal stimulus packages totaled nearly six percent of 2008 nominal GDP. Among the G20 nations, total stimulus was over $2 trillion. This spending generated widespread debate over acceptable levels of national debt (notably, Reinhart and Rogoff, 2010 and Herndon et al., 2013). Austerity measures imposed in several European nations led to criticism and protests from citizens facing economic hardship (IILS, 2011). In mid-2006, the U.S. current account deficit was more than six percent of nominal GDP, the highest in over forty years. A 2004 European Central Bank report labeled "Large and growing U.S. current account deficits" as "...significant risk[s] for global financial stability" (Obstfeld and Rogoff, 2009). Large current account imbalances were not isolated to the United States, however. Japan and Germany saw current account surpluses of almost five and eight percent of GDP, respectively, with a current account deficit of over seven percent of GDP in Australia.

Worldwide financial and capital flows are often asymmetric. While some of these imbalances have diminished since the Great Recession, the U.S. continues to act as a safe haven for financial inflows from around the world. Leading up to the financial crisis, concern over the risks these imbalances posed to the global macroeconomy were well documented by leading economists. Some suggested that a worldwide recession would be unavoidable should these current account imbalances continue (Hausmann and Sturzenegger, 2005). Adding to this issue's complexity is the variety of policy actions and economic circumstances characterizing these economies during this time period. The profile of the current account balance in the U.S. has increased over the last forty years given a three-fold increase in real GDP over this period plus the emergence of the dollar as the world's leading international reserve currency.

In response to the global imbalance dialogue and the search for macroeconomic cause and effect that followed the recent global financial crisis, I investigate the twin deficits question using a panel of eight countries. Data from Australia, Canada, France, Germany, Italy, Japan, the United Kingdom, and the United States are pooled to study the effects of general, non-idiosyncratic shocks – much like those that were felt worldwide from 2007 to 2009 – on government expenditures and the current account. This approach is unique among existing empirical work on this topic. I estimate a vector autoregression (VAR) model using Bayesian methods and an analytical prior to study these shocks in a representative advanced economy. I divide these eight countries into subgroups to examine differences in deficit comovement and shock responses over longer time periods. After controlling for unobserved country-specific effects and the time periods spanning the global financial crisis of 2007 to 2009, we find that increases in government spending do not lead to deterioration of the current account.

The paper proceeds as follows. Section 1.2 discusses aspects of the twin deficits hypothesis studied by the existing literature. Sections 1.3 and 1.4 introduce the model and describe the construction of the panel dataset. Section 1.5 addresses the Bayesian estimation method, prior selection, and presents the results. Finally, Section 1.6 concludes.

1.2 Literature Review

Is the current account a barometer of national economic health? In retrospect, the global imbalances referenced above seem to have been clear warning signs for the financial crisis that followed. At the time, however, some disagreed. Corden (2007) argues that current account surpluses and deficits are simply the result of rational, country-specific spending and saving decisions. Whether they refer to high saving rates in various Asian nations or externally-financed consumption in the U.S. and the U.K., the resulting imbalances reflect market equilibrium. Others have pointed to the fact that the magnitude of gross financial flows between nations dwarf the reported net balances as reason to overlook the current account as an important economic indicator. As Johnson (2009) details, netting inflows and outflows against each other implicitly assumes that the risk characteristics of assets and liabilities are similar, obscuring the fact that these assets and liabilities are not available to cover one another in the event of liquidity shortage or financial instability.

As Obstfeld (2012) outlines, the uncertainty surrounding the significance of current account imbalances highlights the need to track and study their movement. Research from Jordá, Schularick, and Taylor (2011) and Schularick and Taylor (2012) has shown that statistically significant increases in domestic credit issuance foreshadow national and global economic crises. They present empirical evidence suggesting current account deterioration may precede economic slowdowns. Furthermore, current account balances help identify nations that rely on financial inflows to finance spending in excess of domestic saving. These countries may be at higher risk of economic disruption given changes in foreign investment at the onset of global economic turmoil.

Previous empirical work on twin deficits consists primarily of time series data analyses for individual nations. Effects of non-country specific shocks or interactions between nations have not been considered. Bernheim (1988) uses OECD data from 1960 to 1984 to show that trade deficits in the U.S., Canada, U.K., Mexico, and West Germany all increase in response to rising government budget deficits. Kearney and Monadjemi (1990) use a VAR model with quarterly data from the 1970s and 1980s to suggest that the twin deficits relationship breaks down in the long run. Leachman and Francis (2002) show that relationships between the budget balance and the trade balance in the U.S. may be time specific. Using data stretching back to the Bretton Woods fixed exchange rate system, the authors dismiss the twin deficits hypothesis in the post-WWII period of 1948 to 1973, but find weak evidence of twin deficits in the post-Bretton Woods period from 1974 to 1992. Mohammadi (2004) uses panel data to examine the effects of increases in government spending financed via borrowing versus via tax increases, finding evidence in favor of twin deficits.

To investigate this paradox, Kim and Roubini (2008) (hereafter K&R) use a VAR model to identify fiscal deficit shocks and their effects on real output, the current account, and the real exchange rate in the United States. They argue that "twin divergence" is the more accurate description of what has been seen in the United States since 1973. Except for short periods in the 1980s and the early 2000s, the U.S. government budget deficit does not move in tandem with the U.S. current account deficit. They find that shocks to the primary government budget deficit lead to short-term improvement in the current account and an increase in the real exchange rate (domestic currency depreciation). The latter result also contradicts the twin deficits hypothesis since, in theory, domestic currency appreciation should follow a rise in the primary government budget deficit. Numerous other studies, including Ratha (2012), Makin and Narayan (2013), El-Baz (2014), and Baharumshah, Lau, and Khalid (2004), also investigate twin deficits in individual countries. These studies find mixed evidence for and against the twin deficits hypothesis.

1.3 Panel VAR Model

I use a vector autoregression (VAR) model (Sims, 1980) with quarterly data to estimate the effects of government spending shocks on real output, the current account, the real interest rate, and the real exchange rate. The definitions of these five variables are listed in Table 1.1.

The data are from the OECD Main Economic Indicators database and were obtained via *Federal Reserve Economic Data (FRED)*. RGDP is constructed using seasonally adjusted nominal GDP along with each nation's GDP price deflator to obtain real GDP in year 2000 units of national currency. Logs are taken and the series is differenced to form growth rates. GOV is calculated by dividing nominal government expenditures in each quarter by that quarter's nominal seasonally adjusted GDP figure.[1] CUR is equal to the current account balance divided by nominal GDP. The real interest rate RIR is calculated by taking each

[1]Nominal government expenditure data for Australia is not available. Instead, real government expenditure data is adjusted to year 2000 Australian dollars.

country's three-month interbank rate and subtracting the respective GDP price deflator annual rate of change. Finally, the real exchange rate variable (RER) is constructed by taking the log of the indexed OECD data series. I take the negative of this result to obtain the traditional increase–depreciation and decrease–appreciation convention. Figures 1.1 and 1.2 show GOV and CUR from 1995 to 2012 for each nation in the panel. Figure 1.3 shows similar data for a subset of four nations (Canada, Germany, the United Kingdom, and the United States) over a longer sample from 1982 to 2012. Summary statistics for the entire dataset are shown in Table 1.2.

My panel dataset and the method in which it is constructed is unique. While many applied microeconomics panels have large cross section and small time series dimensions, mine is the opposite: the time series dimension is larger than the cross section. This is typical of macroeconomic panels that involve multicountry analyses (for precedents, see Canova, Ciccarelli, and Ortega (2012); Canova and Ciccarelli (2009); and Ciccarelli, Maddaloni, and Peydró (2013)). These differences are primarily related to data availability constraints, since there are few nations for which quarterly data is available for all five variables I am using. To avoid the pitfalls of applying asymptotic theory relying on on large N values to my panel, I use Bayesian methods for estimation. The balanced panel consists of sixty-seven ($T = 67$) quarterly observations for each country, with data beginning in the second quarter of 1996 (2Q96) and ending in the fourth quarter of 2012 (4Q12). As referenced above, the countries selected are those that make up the G7 plus Australia ($N = 8$). As in K&R, the VAR contains four lags of each of the variables in Table 1.1. To capture the representative effects of shocks to government spending, I pool the data into one group and stack the data for individual nations to generate an overall sample of 536 observations. To account for unobserved effects, dummy variables are included to capture idiosyncrasies pertaining to individual nations and to the quarters spanning the global financial crisis. This structure is formalized below.

1.4 Constructing an International Panel

Before aggregating across countries and time periods, the model is given by (4.1) below. Column vector y_{it} contains current (period t) values of RGDP, GOV, CUR, RIR, and RER for nation i. Column vector x_{it} contains a constant term plus four quarters of lagged values of each of these five variables in the same nation i. Additionally, dummy variables c_i and v_t are included to capture unobserved effects specific to individual countries and the quarters spanning the global financial crisis. Thus, for each nation $i = 1, \ldots, N$, and for each time period $t = 1, \ldots, T$:

$$y'_{it} = x'_{it}\beta + c_i + v_t + u'_{it}. \tag{1.1}$$

To pool these data into one large group, first I aggregate y'_{it}, x'_{it}, and u'_{it} over t into $\mathbf{y_i}$, $\mathbf{x_i}$, and $\mathbf{u_i}$, each of which contains all observations, across time, for a country i:

$$\mathbf{y}_i = \begin{bmatrix} y'_{i1} \\ y'_{i2} \\ \vdots \\ y'_{iT} \end{bmatrix}, \quad \mathbf{x}_i = \begin{bmatrix} x'_{i1} \\ x'_{i2} \\ \vdots \\ x'_{iT} \end{bmatrix}, \quad \text{and} \quad \mathbf{u}_i = \begin{bmatrix} u'_{i1} \\ u'_{i2} \\ \vdots \\ u'_{iT} \end{bmatrix}.$$

When aggregating across countries, the dummy variable v_t is configured into a $T \times 7$ matrix $\mathbf{v_i}$ that accounts for unobserved financial crisis effects in country i. Dummy variables are added for each quarter from 4Q07 to 2Q09 (seven quarters total) such that

$$\mathbf{v_i} = \begin{bmatrix} 0_{50\times 7} \\ \hline I_7 \\ \hline 0_{10\times 7} \end{bmatrix}$$

where I_7 is the 7×7 square identity matrix and $0_{50\times 7}$ and $0_{10\times 7}$ are matrices of zeros. Thus, (4.1) becomes

$$\mathbf{y}_i = \mathbf{x}_i\beta + \mathbf{v_i}\gamma + \mathbf{u}_i. \tag{1.2}$$

Finally, I construct a linear model that combines each individual country's data into a single sample. This procedure pools the data in the panel into one representative economy and eliminates the need to estimate coefficients for each individual nation. Each country's $\mathbf{y}_i$, $\mathbf{x}_i$, and $\mathbf{u}_i$ are stacked to form $\mathbf{Y}$, $\mathbf{X}$, and $\mathbf{U}$ that contain all current and lagged values for RGDP, GOV, CUR, RIR, and RER. $\mathbf{X}$ includes a column of ones for a constant term.

$$\mathbf{Y} = \begin{bmatrix} \mathbf{y}_1 \\ \mathbf{y}_2 \\ \vdots \\ \mathbf{y}_N \end{bmatrix}, \quad \mathbf{X} = \begin{bmatrix} \mathbf{x}_1 \\ \mathbf{x}_2 \\ \vdots \\ \mathbf{x}_N \end{bmatrix}, \quad \text{and} \quad \mathbf{U} = \begin{bmatrix} \mathbf{u}_1 \\ \mathbf{u}_2 \\ \vdots \\ \mathbf{u}_N \end{bmatrix}.$$

Matrices $\mathbf{D}$ and $\mathbf{V}$ contain the country and financial crisis dummy variables, respectively, to be used as additional regressors. Since the model contains a constant, $N-1$ country dummy variables (one for each non-U.S. nation in the panel) are added. Furthermore, for each country, the $T-1$ financial crisis dummy variables described above are also included:

$$\mathbf{D} = \begin{bmatrix} D_1 \\ \hline D_2 \end{bmatrix} \quad \text{and} \quad \mathbf{V} = \begin{bmatrix} \mathbf{v_1} \\ \vdots \\ \mathbf{v_N} \end{bmatrix},$$

where $D_1 = I_{N-1} \otimes \iota_T$ has dimension $(N-1)T \times (N-1)$, D_2 is a matrix of zeros with dimension $T \times (N-1)$, and I_{N-1} is the $N-1$ square identity matrix. Combining the regressors such that

$$\tilde{\mathbf{X}} = \left[\begin{array}{c|c|c} \mathbf{X} & \mathbf{D} & \mathbf{V} \end{array}\right],$$

the global model can be expressed more compactly as

$$\mathbf{Y} = \tilde{\mathbf{X}}\mathbf{B} + \mathbf{U}. \tag{1.3}$$

With the panel consisting of $N = 8$ nations, the matrix $\mathbf{B}$ consists of five columns of coefficients, one for each variable in $\mathbf{Y}$. The rows of $\mathbf{B}$ contain coefficients for each of four lags for each of the five variables, a constant term, the coefficients on the dummy variables for the $N - 1$ nations, and the coefficients on the dummy variables for the quarters during the financial crisis.

This dataset provides insight into the effects of government spending in an advanced open economy. My analysis takes the point of view of a representative economy as opposed to describing country-specific responses that have been the primary focus of previous empirical research concerning twin deficits. This representative economy is formed by aggregating data from eight nations as detailed above. I describe my results in the next section.

1.5 Estimation Procedure and Results

Out of consideration for the lack of clear consensus regarding the accuracy of the twin deficits hypothesis, the model is estimated with a Normal-Wishart prior. The Normal-Wishart prior is a conjugate prior for the multivariate normal distribution. This prior is computationally simple and permits direct sampling without use of approximation techniques like Gibbs sampling. The Normal-Wishart prior extends the Minnesota prior of Litterman (1986) by drawing the inverse of the variance-covariance matrix of the residuals from a Wishart distribution rather than assuming it is fixed. To describe the Normal-Wishart prior, I follow the notation used in Koop and Korobilis (2010).[2] The Normal-Wishart prior has the form

[2]Code adapted from `http://personal.strath.ac.uk/gary.koop/bayes_matlab_code_by_koop_and_korobilis.html` is used to estimate the model.

$$\alpha|\Sigma \sim \mathcal{N}(\underline{\alpha}, \Sigma \otimes \underline{V}) \tag{1.4}$$

where

$$\Sigma^{-1} \sim W(\underline{S}^{-1}, \underline{\nu}). \tag{1.5}$$

where $\mathcal{N}$ and W denote the multivariate normal and Wishart distributions, respectively. In (5.1), $\alpha = vec(\mathbf{B})$ from (4.3) above. The values of hyperparameters $\underline{\alpha}, \underline{V}, \underline{S}$, and $\underline{\nu}$ are chosen as follows. Litterman (1986) suggests setting prior coefficient means $\underline{\alpha}$ based on the unit of measurement of each variable. Variables measured in log-levels (i.e., GOV, CUR, RIR, and RER) are assumed to follow a random walk pattern and have initial values of one on the first own lag and zero elsewhere. For RGDP, measured as a growth rate, the prior mean on all coefficients is zero. Also set to zero are the prior means of non-lagged variables (the constant, plus the time and country dummies). Exogenous foreign demand for domestic goods and services, a multitude of policy choices affecting fiscal spending and interest rates, and a lack of empirically verified theories of exchange rate determination contribute to the choices for these hyperparameters.

I construct the prior variance-covariance matrix $\underline{V}$ according to Karlsson (2013). Hyperparameters $\pi = (\pi_1, \pi_3, \pi_4)$ control overall prior tightness, the speed at which lagged values shrink to zero, and the probability weight placed on deterministic variables (dummies and constants), respectively.[3] Benchmark parameter settings for π are (.2, 2, 100,000). These values are derived from initial values suggested in Canova (2007). Additional hyperparameter settings assigning various probability weight to prior and sample information are also tested. I find that they have no material effects on the responses of the variables to shocks to government spending.

$\underline{S}$ and $\underline{\nu}$ define the distribution from which the variance-covariance matrix of the residuals, Σ, is drawn. $\underline{S}$ is initially defined to be a diagonal matrix with entries from univariate

[3] In my implementation of the Normal-Wishart prior, the value π_2 controls the relative tightness on own lagged variables versus lags of other variables. It is normalized to one.

autoregressive models for each independent variable. In these models, current period values are regressed on four quarters of own lagged values and a constant. The estimates of the error variance form the entries of $\underline{S}$. $\underline{\nu}$ controls the degrees of freedom of the prior. In this model, $\underline{\nu} = 7$.

Values for the coefficients in (4.3) are drawn directly from the multivariate normal posterior distribution given by

$$\alpha|\Sigma \sim \mathcal{N}(\bar{\alpha}, \Sigma \otimes \bar{V}) \tag{1.6}$$

where

$$\Sigma^{-1}|y \sim W(\bar{S}^{-1}, \bar{\nu}). \tag{1.7}$$

with $\bar{V}, \bar{\mathbf{B}}, \bar{S}$, and $\bar{\nu}$ given by

$$\begin{aligned} \bar{V} &= [\underline{V}^{-1} + \tilde{\mathbf{X}}'\tilde{\mathbf{X}}]^{-1} \\ \bar{\mathbf{B}} &= \bar{V}[\underline{V}^{-1}\underline{\mathbf{B}} + \tilde{\mathbf{X}}'\tilde{\mathbf{X}}\hat{\mathbf{B}}] \\ \bar{S} &= S + \underline{S} + \hat{\mathbf{B}}'\tilde{\mathbf{X}}'\tilde{\mathbf{X}}\hat{\mathbf{B}} + \underline{\mathbf{B}}'\underline{V}^{-1}\underline{\mathbf{B}} - \bar{\mathbf{B}}'\underline{V}^{-1} + (\tilde{\mathbf{X}}'\tilde{\mathbf{X}})\bar{\mathbf{B}} \\ \bar{\nu} &= NT + \underline{\nu}. \end{aligned} \tag{1.8}$$

In (5.3), $\bar{\alpha} = vec(\bar{\mathbf{B}})$. In (5.5), $\underline{\mathbf{B}}$ is formed by transforming $\underline{\alpha}$ from a vector back into a matrix (i.e. back into the shape of $\mathbf{B}$), and $S = (\mathbf{Y} - \tilde{\mathbf{X}}\hat{\mathbf{B}})'(\mathbf{Y} - \tilde{\mathbf{X}}\hat{\mathbf{B}})$, where $\hat{\mathbf{B}}$ is the OLS estimate of $\mathbf{B}$.[4]

The structural shocks of the model are identified using a Choleski decomposition on Σ. Impulse responses in Figures 1.4 and 1.5 display the median response from 10,000 posterior draws. Also shown in these graphs are ninety percent confidence bands generated by displaying the 5th and 95th percentiles of impulse responses from 10,000 draws. I focus on several groups of countries: the complete panel with all G7 nations plus Australia; the complete panel excluding the United States; the six North American and European nations of

[4]Koop and Korobilis (2010), Kadiyala and Karlsson (1997), and Karlsson (2013) provide complete mathematical derivations of the items in (4.5), (4.6), (4.7).

Canada, France, Germany, Italy, the United Kingdom, and the United States; and a smaller panel including Canada, Germany, the United Kingdom, and the United States for which longer-term data is available. I am most interested in the responses of CUR and RER to shocks to GOV, as these were the main predictions of the twin deficits hypothesis.

The impulse responses for the full panel of eight countries (row one of Figure 1.4) indicate that the current account does not change significantly following a one percent increase in government spending as a percentage of GDP. Real interest rates rise significantly before retreating after approximately three quarters. The real exchange rate depreciates slightly, which is consistent with the findings of K&R. Output decreases somewhat, but the majority of the response of real GDP to shocks to GOV is not significant. In the non-U.S. panel (row two of Figure 1.4), the responses of each variable are nearly identical to the full panel. For the non-APAC panel (row three of Figure 1.4), the current account deteriorates slightly with real interest rates rising significantly and no change in the exchange rate. In row four of Figure 1.4, the responses of a subset of four nations – Canada, Germany, the U.K., and the U.S. – differ from the larger panels in rows one through three. Output and real interest rates decrease significantly, a slight improvement in the current account balance.

My results do not show the current account deteriorating significantly in response to increases in government spending, which I use as a proxy for increases in the government budget deficit.[5] The slight depreciation in exchange rates seen in the full and non-U.S. panels is small and lasts less than two quarters. In each panel, real GDP growth falls initially. This may be the result of the government spending shock "crowding out" investment before interest rates decline as the shock fades. The negative movement in output on shock impact may also be the product of consumers reducing consumption and supplying more labor in response to increases in government spending as predicted by the neoclassical model.

To check the sensitivity of the results to the time period we have selected, I extend the

[5] While is is not a perfect solution to the lack of quarterly government budget deficit data, it is reasonable to assume that these countries would finance additional spending via borrowing given that each country, with the slight exception of Germany, has consistently run annual budget deficits.

analysis to include much of the 1980s. Figure 1.5 shows impulse responses from a four-country panel (Canada, Germany, the U.K., and the U.S.) over various time periods covering the early 1980s to 2012. These panels are constructed in the same manner described in Sections 1.3 and 1.4.

Row one of Figure 1.5 includes impulse responses generated from data spanning the second quarter of 1982 to the second quarter of 2012. This longer sample shows a slight decline in the current account following a shock to government spending. The real interest rate decreases briefly after approximately four quarters. There is no immediate movement in exchange rates, though some depreciation is seen after several years. In rows two and three of Figures 1.5 I consider two mutually exclusive sub-periods of the longer period in row one. In neither case do I find significant current account deterioration or exchange rate appreciation following increases in government spending. In each of these three panels, real output decreases significantly on shock impact. Overall, these responses do not provide evidence confirming the twin deficits hypothesis. My results have shown this for a variety of country groupings and time periods.

1.6 Conclusion

This work has shown how pooling data from multiple nations into a macroeconomic panel provides empirical international evidence that the primary claim of the twin deficits hypothesis – that increases in the government budget deficit lead to increases in the current account deficit – is not supported by an analysis that uses panel data to create a representative advanced economy. In several cases, the first chapter of the twin deficits story is confirmed: real interest rates rise significantly following positive shocks to government expenditures. However, my results do not consistently show downward movement in the current account or appreciation of real exchange rates.

The goal of this paper has been to improve knowledge of the relationship between government spending and the current account balance. As alluded to in Section 1.2, the current account plays an important role in portending future economic trouble. Forthcoming research will focus on expanding these findings to generate new conclusions, focusing specifically on the use of panel data to investigate the effects of government spending. Once these basic macroeconomic connections are better understood, economists will be better equipped to recognize looming economic problems. I expect that these questions will continue to be leading objects of interest for policymakers in both the near term and the long term.

Variable	Description
RGDP	*Growth rate of real GDP; GDP measured in year 2000 local currency*
GOV	*Government expenditures as a percentage of nominal GDP*
CUR	*Current account as a percentage of nominal GDP*
RIR	*3-month interbank rates, adjusted for inflation*
RER	*Log of real effective exchange rate based on manufacturing consumer price index*

Table 1.1: Variable definitions

Variable	**Obs.**	**Mean**	**Std. Dev.**	**Min**	**Max**
RGDP	536	0.19%	0.34%	-1.83%	1.22%
GOV	536	18.93%	2.96%	12.01%	25.04%
CUR	536	-0.58%	3.23%	-7.19%	7.77%
RIR	536	1.51%	2.59%	-7.64%	15.36%
RIR	536	0.01	0.14	-0.28	0.42

Table 1.2: Summary statistics for full panel

Figure 1.1: First half of plots of GOV and CUR, 1995 to 2012
GOV (solid line) and CUR (dashed line) are each measured as a percentage of nominal GDP.

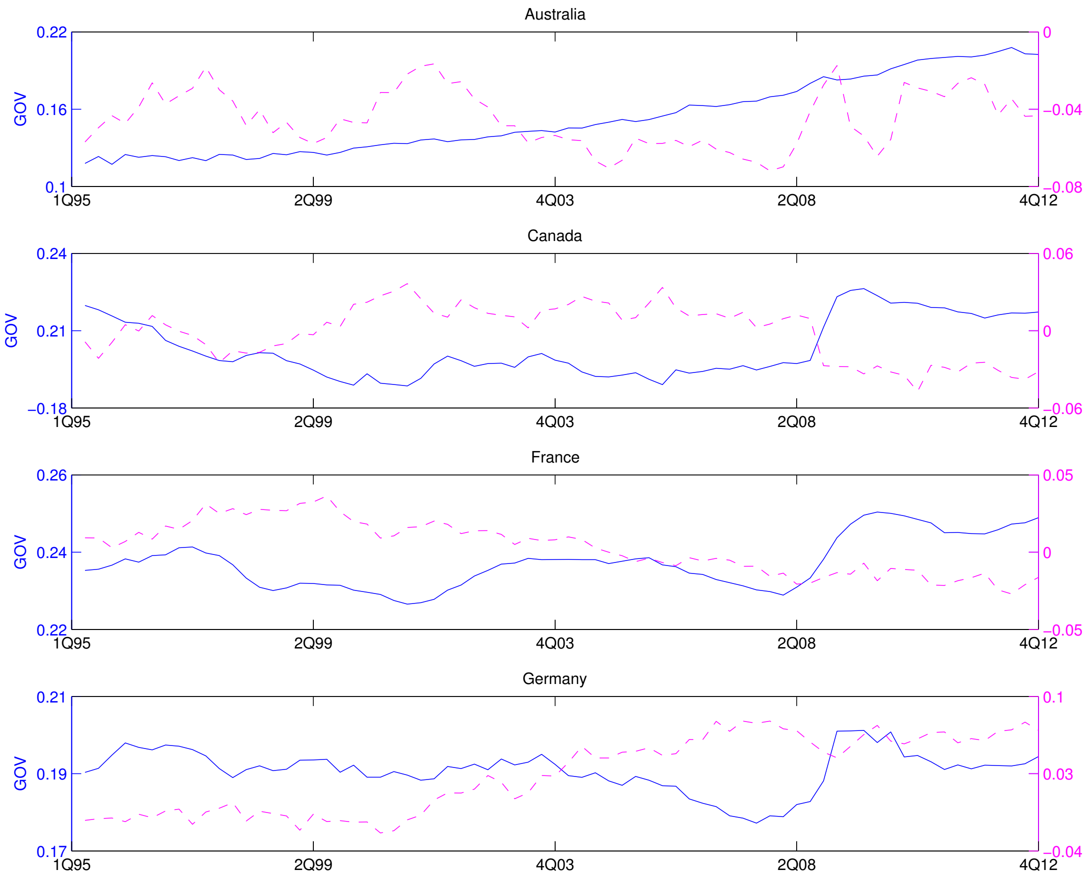


Figure 1.2: Second half of GOV and CUR, 1995 to 2012
GOV (solid line) and CUR (dashed line) are each measured as a percentage of nominal GDP.

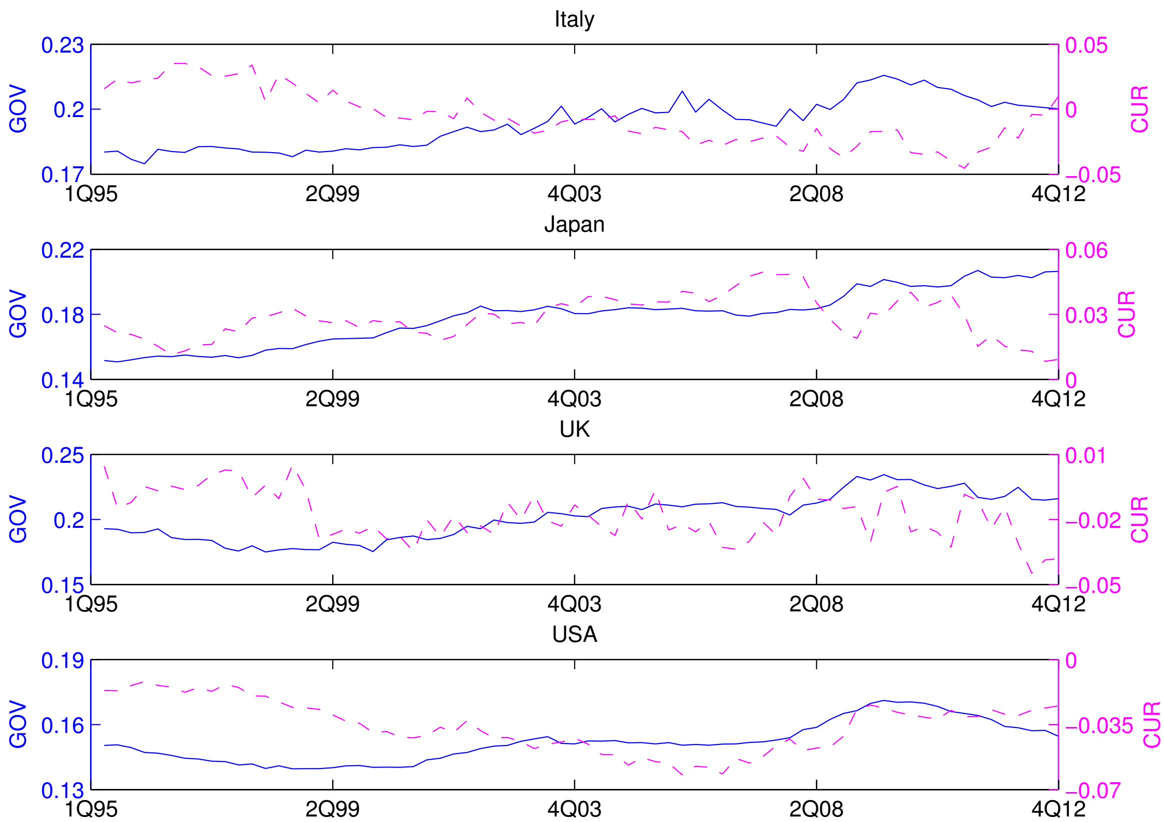
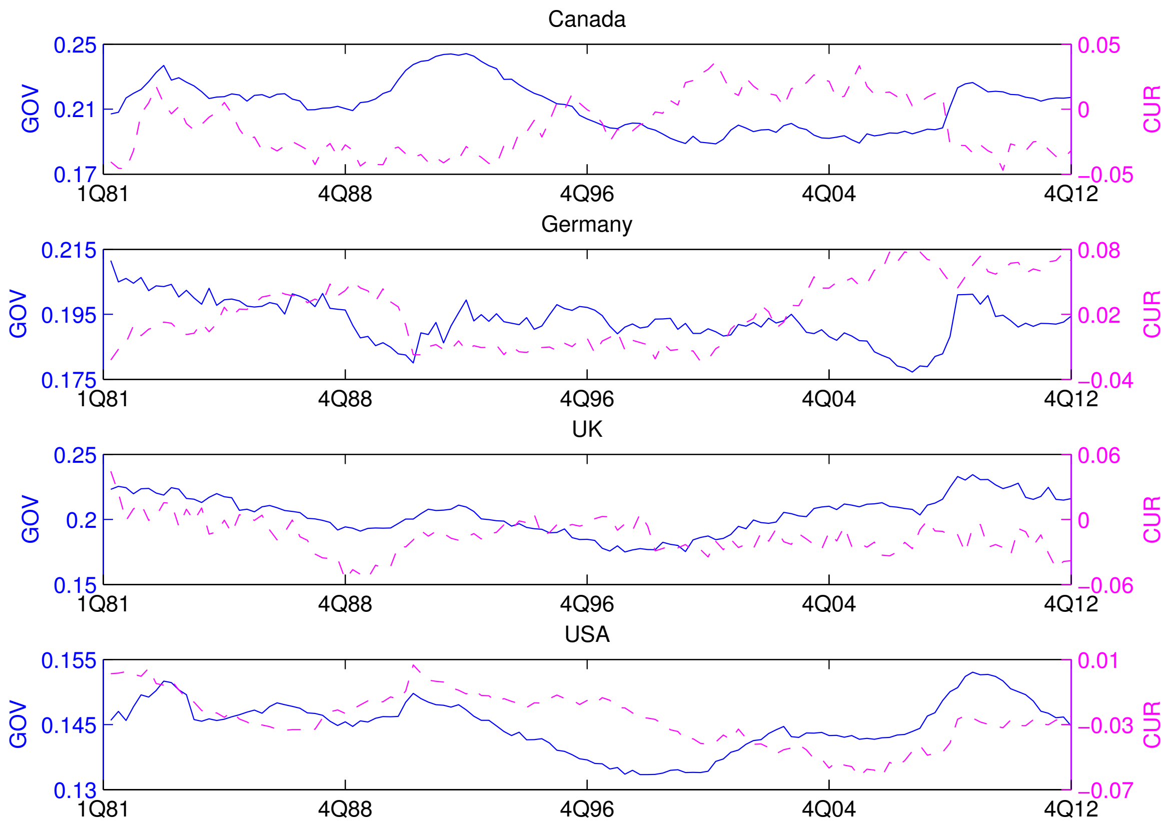


Figure 1.3: Plots of GOV and CUR, 1981 to 2012
GOV (solid line) and CUR (dashed line) are each measured as a percentage of nominal GDP.



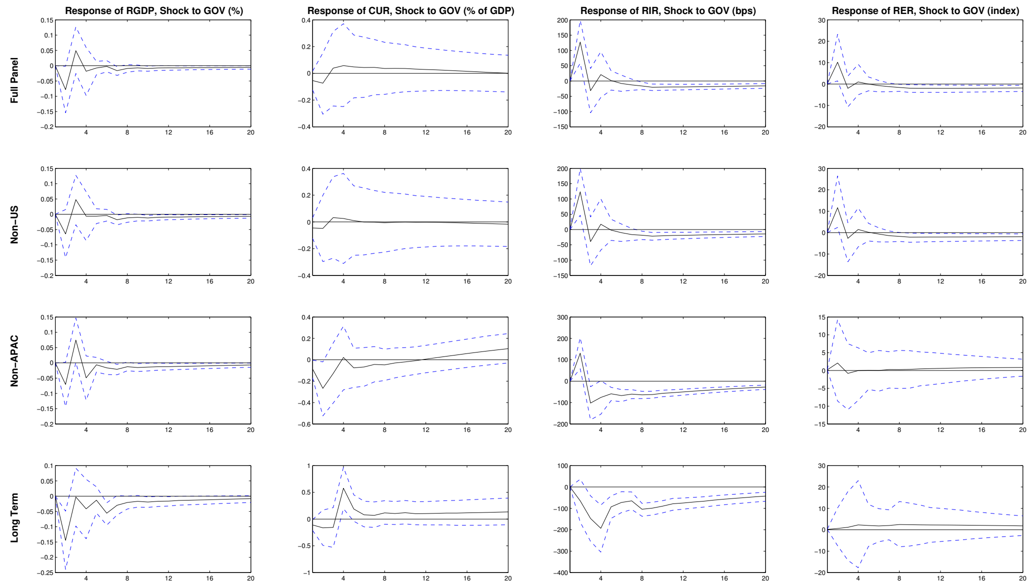


Figure 1.4: Impulse responses of 1% shock to GOV for various panels, 2Q96 to 4Q12
First row includes data from all eight nations; second row, data from all non-U.S. nations; third row, data from Canada, France, Germany, Italy, U.K., U.S.; fourth row, data from Canada, Germany, U.K., U.S. Dashed lines indicate ninety percent confidence bands.

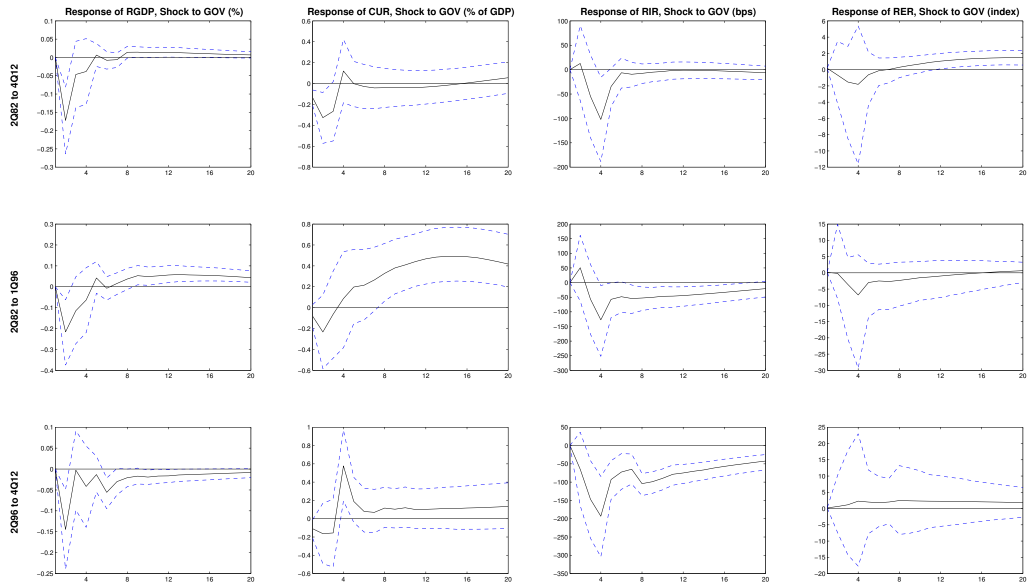


Figure 1.5: Impulse responses of 1% shock to GOV for four-country panel, various periods Four-country panels includes data from Canada, Germany, U.K., U.S. First row includes data from the period 2Q82 to 4Q12; second row, 2Q82 to 1Q96, third row, 2Q96 to 4Q12. Dashed lines indicate ninety percent confidence bands.

Chapter 2

Identifying the Effects of Government Spending Shocks in a Macroeconomic Panel

2.1 A Challenging Macroeconomic Question

Confusion surrounding the specific effects of changes in government spending on consumption, wages, output, and interest rates is costly for researchers and policymakers. As economies move along the path of the business cycle, fiscal spending is a tool for central banks and governments to smooth disturbances and sustain economic progress. But uncertainty surrounding the correct scale and timing of proposed stimulus packages can limit their welfare effects. Government stimulus packages enacted worldwide in response to the global financial crisis of the late 2000s were designed to mitigate recessions and widespread job losses that affected many countries. In the United States, the American Reinvestment and Recovery Act (ARRA) earmarked $840 billion to be spent over a ten-year period lowering taxes, increasing unemployment benefits, and funding construction and infrastructure projects (Congressional Budget Office, 2015). Australia, Canada, Italy, France, and the

United Kingdom each used combinations of infrastructure spending, tax relief, and labor market interventions to jump-start their economies during the post-crisis years (IILS, 2011).

With the multitude of policy actions taken during this period – from changes in interest rates to troubled asset relief – it is very difficult to quantify the orthogonal effects of these stimulus packages. Ideally, economic models studying the effects of fiscal spending should isolate variation that is independent of business cycle dynamics. However, truly exogenous changes in government spending are difficult to extract given the noisy nature of macroeconomic data, the large potential for simultaneity, and the challenges in addressing lags and timing issues associated with fiscal policy.

In the absence of controlled experiments, macroeconomists can use creative research design to reduce endogeneity in empirical analyses. One such method, the narrative approach, is a non-statistical technique that identifies variation in time series data by reviewing the historical record. By creating variables derived from media reports, policy records, and official data sources describing changes outside the business cycle, economists can test how macroeconomic variables of interest respond in the months and quarters following shocks. This procedure stands in contrast to the identification approaches of structural VAR models that apply shocks to total government spending and find contrasting effects on consumption and wages relative to studies identifying government spending shocks using military buildups. This dichotomy has implications for governments and policymakers, and for economists whose theoretical models are evaluated in comparison with economic data.

In the analysis that follows I utilize the narrative approach to develop a measure of government spending changes in a panel of mature, developed nations. From articles appearing in national newspapers, I extract the dates, amounts, and spending horizons of changes in military and civil defense expenditures related to military buildups, acts of terrorism at home and abroad, and other geopolitical events over the period since the September 11 attacks in the United States. I have identified this period as one in which frequent and meaningful changes in defense and civil defense spending levels were motivated by factors independent of the global business cycle. This period provides a unique experimental setting where I can more clearly assess causality between macroeconomic variables. I record these spending amounts in the form of a defense and civil defense news variable, *defnews*, and show that shocks to *defnews* correctly identify changes in total government spending. I use this variable in a panel VAR model with output, interest rates, consumption, unemployment, and wages in six advanced economies to develop a new set of stylized facts describing the effects of changes in government spending that are disconnected from the business cycle.

By gathering data from multiple nations, I am able to increase the dimensions of my dataset in spite of a sample period that is short relative to empirical studies reviewing several decades of data. As discussed in Canova and Ciccarelli (2009, 2013), panel, or multicountry, VARs combine time series data from a collection of nations, markets, or sectors. Panel VARs have previously been used to examine business cycle dynamics in groups of nations (e.g., Canova and Ciccarelli (2012)), to study the spillover effects of government spending shocks across borders (Beetsma and Giuliodori (2011)), and to investigate the transmission of financial shocks between countries in an optimum currency area (Ciccarelli, Maddaloni, and Peydró (2013)). My multicountry VAR model provides estimates of the representative effects of changes to government spending in an advanced economy. The model generates these estimates while also accounting for factors specific to each country that could influence these representative estimates.

The structure of the panel, which has a short cross section and a moderate time series length, precludes me from using certain classical methods that rely on larger dimensions to sustain the asymptotics required for efficient and unbiased inference. For this reason, and because of the benefits provided by shrinkage in a model with over 200 unknown parameters, I estimate my model using Bayesian methods. The shrinkage provided by the prior distribution enables me to more easily manage the large number of parameters in my model and reduce much of the uncertainty in the responses of the dependent variables. In conjunction with my uniquely constructed *defnews* variable, this approach sharpens the conclusions I make regarding the effects of changes in government spending.

Similar to Ramey (2011), my estimates of the effects of increases in government spending on consumption, unemployment, and the real wage support the results of the neoclassical model. The neoclassical model predicts that households reduce consumption and supply more labor in response to increases in government expenditures financed by lump-sum taxes. In the short term, this reduces the equilibrium real wage and increases the marginal product of capital. Interest rates rise, fueling an increase in investment; capital accumulates and the real wage returns to its steady state level. The results from my panel VAR model show that consumption and unemployment each fall in response to an increase in government spending. These results stand in contrast to New Keynesian models predicting increases in consumption and the real wage following shocks to government purchases. I provide new evidence explaining the macroeconomic responses in advanced economies to defense and civil defense spending events that have become increasingly common over the last fifteen years.

The paper proceeds as follows. Section 2.2 reviews the relevant pieces of the existing literature. Section 2.3 describes my Ramey (2011)-inspired instrument for defense spending and the data collection process that generated it. Section 2.4 describes the rest of the data, the model, and the estimation procedure. Section 2.5 presents the results and Section 2.6 provides concluding remarks.

2.2 Review of Precedent Literature

2.2.1 The Theoretical and Empirical Effects of Government Spending

According to neoclassical theory, an increase in government spending financed by lump sum taxes triggers a negative wealth effect that lowers consumption and precipitates an increase in agents' labor supplied. This lowers the equilibrium real wage and increases the marginal product of capital. The real interest rate rises, capital accumulates, and the marginal product of capital falls, with the real wage declining to its original level. Hall (1980), Barro (1981), Aiyagari, Christiano, and Eichenbaum (1992), and Baxter and King (1993) each examine the effects of changes in government spending in a neoclassical framework. Hall (1980) and Barro (1981) find that hours worked and output are expected to increase under either temporary or permanent increases in government spending. Aiyagari, Christiano, and Eichenbaum (1992) and Baxter and King (1993) find that the response of output is greater when the government spending increase is permanent. In contrast, Blanchard and Perotti (2002) use a structural VAR model and U.S. data to show that increasing government spending leads to increases in output and consumption. Perotti (2007) utilizes a structural VAR approach with long-run annual data and finds evidence that the response of consumption to government spending shocks can be positive. Galí, López-Salido, and Vallés (2007) use a similar approach, concluding that consumption and the real wage increase following shocks to government spending. They propose a New Keynesian model with sticky prices and rule-of-thumb consumers to justify these results.

In these traditional VAR models (the aforementioned Blanchard and Perotti (2002), Perotti (2007), and Galí, López-Salido, and Vallés (2007)), assumptions made regarding the ordering of variables – specifically, that government spending is exogenous to all other variables in the system – is ubiquitous but unproven. If output determines taxes, and subsequently the level of government expenditure, including consumption, taxes, and various types of investment as additional variables in these VARs leads to questions about the validity of the ordering.

Exploiting military buildups as sources of non-business cycle variation in government spending is used throughout the literature to address the endogeneity concerns inherent in traditional VAR analyses. This strategy relies on two key factors. First, changes in military spending are unlikely to be correlated with the business cycle. As Hall (2009) writes, "military spending does not respond to forces determining GDP or consumption... but only to geopolitical events." He also asserts: "the most direct way to measure the government purchases multiplier is to exploit large and arguably exogenous fluctuations in military spending." Using military spending as the lens through which to view shocks to government spending reduces researchers' exposure to endogeneity criticisms.

Second, the types of consumption undertaken by governments for defense purposes is likely to be distinct from the types of consumption undertaken by households. It is reasonable to expect that the goods purchased with defense and civil defense funds are unlikely to affect contemporaneous labor-leisure tradeoffs of households. As Ramey (2011) notes, aggregate government spending includes expenditures related to welfare programs, education, and institutions. Spending on these items may influence households' choices and muddle estimates drawn from models using total government expenditures as the source of variation.

Ramey and Shapiro (1998) construct a series that uses dummy variables to identify large military buildups in the U.S. at three key dates: the Korean War, the Vietnam War, and the Soviet invasion of Afghanistan. Estimates from an autoregressive model indicate increases in output and declines in consumption and real wages consistent with a neoclassical model. Edelberg, Eichenbaum, and Fisher (1999) present an extension of Ramey and Shapiro in which shocks to government spending elicit increases in output and employment but decreases in consumption and the real wage. Ramey (2011) shows that shocks to government spending in traditional VAR models can be predicted by the "war dates" shocks of Ramey and Shapiro, and that differences in the responses of consumption and the real wage between the two approaches are consequences of timing. Due to lengthy processes related to procurement and awarding of contracts, delays between the dates of defense spending announcements and the actual increases in government spending they generate lead traditional VAR models to overlook initial declines in consumption and wages following shocks to government spending. Ramey shows that shocks identified by the Ramey-Shapiro war dates approach Granger-cause traditional VAR shocks, and that delaying the Ramey-Shapiro war dates by several periods generates impulse responses matching those of the traditional VAR models. She enriches the Ramey-Shapiro war dates by constructing a "defense news" variable that measures the present discounted value of defense spending as a percentage of lagged nominal GDP. Shocks to this defense news variable produce declines in consumption and the real wage consistent with Ramey and Shapiro (1998).

2.2.2 The Narrative Approach and Natural Experiments in Economics

Both Ramey and Shapiro (1998) and Ramey (2011) use variables constructed from newspaper and media reports of changes in military spending to construct new measures of government spending. The procedure of extracting relevant information from the historical record, termed the narrative approach, first appeared in Friedman and Schwartz (1963) and has been used more recently by Romer and Romer (1989) to identify monetary policy shocks from meeting minutes and policy records of the Federal Reserve Board of Governors and the Federal Open Market Committee. Romer and Romer (2010) review economic and Congressional reports to identify changes in tax legislation and the resulting effects on output. Romer and Romer (2015) examine historical economic reports prepared by the Organisation for Economic Co-operation and Development (OECD) to construct an index of financial distress that measures the length and severity of financial crises. The variable they construct captures changes in the cost of credit in crisis-stricken nations and allows the authors to quantify the effects of financial crises – and their intensity – on GDP and aggregate economic activity.

As I have described above, applying this technique to identification of government spending shocks is a creative way to tease out variation in macroeconomic time series data that cannot easily be captured by statistical techniques. By focusing on civil defense spending in addition to military buildups, I can study my research question in a natural experiment setting that is generally elusive in social science research. As defined in Fuchs-Schuendeln and Hassan (2015), natural experiments "provide observable, quasi-random variation in treatment subject to a plausible identifying assumption." This definition applies to the nature of changes in civil defense spending: they are highly likely to be the result of geopolitical events (i.e., quasi-random) and uncorrelated with the optimizing decisions of households (a plausible identifying assumption).

2.3 A Richer Measure of Changes in Government Spending

The objective of my unique and extensive data collection process is to find and record all new announcements of changes to military and civil defense spending using the narrative approach. These announcements are made, depending on the country and the legislative process, by prime ministers, presidents, finance ministers, defense ministers, cabinet departments, and legislative bodies. These announcements capture both increases and decreases in expenditures related to emergency situations, acts of terrorism at home and abroad, reinforcements to existing military or civil defense operations, new departments and initiatives, and new capital projects for weapons, buildings, or aircraft. I search databases of media sources, primarily newspapers, to gather data to construct a variable that is the foundation of my analysis. I believe that this variable identifies quasi-exogenous changes to government spending because it captures significant changes in government expenditures that are, for the most part, a response to geopolitical events that occur independently of the business cycle. This differentiates my analysis from the works described above that rely on models identifying what is likely to be endogenous variation in government spending.

The variable I create, *defnews*, is similar to the defense news variable used in Ramey (2011). The differences between what Ramey creates in her U.S.-only dataset and what I construct for a panel of nations are twofold. First, the sample periods are different. My sample period includes only the fifty-four quarters from the third quarter of 2001 to the fourth quarter of 2014. I have identified this period as one in which homeland security became a large priority for the nations in my panel. In response to the 9/11 attacks, French

President Jacques Chirac remarked, "we must fight against terrorism by all means." United Kingdom Prime Minister Tony Blair said, "mass terrorism is the new evil in our world," and that "[this is a battle] between the free and democratic world and terrorism." The 9/11 commission report released in 2004 by the U.S. government acknowledged that "countering terrorism has become, beyond any doubt, the top national security priority for the United States." According to the Human Rights Watch, more than 140 countries, including all six in the panel, passed counterterrorism laws since September 11, 2001. Focusing on the variation in spending levels that took place during a period in which numerous resources were mobilized to protect populations, military personnel, and national borders from the burgeoning threat of terrorism provides a natural experiment setting in which government spending levels changed independently of the state of the global economy.

Second, my *defnews* variable captures changes in civil defense spending in addition to purely military and defense spending. These civil defense expenditures include airport security operations, intelligence services, bioterrorism prevention measures, police resources, and other programs designed to protect borders and civilians. Although the amounts of civil defense spending announcements are often lower than the military spending announcements, their costs still range into the billions of units of national currency. I believe that they complement the exclusively military data points and add a degree of depth to my *defnews* variable above that of Ramey.

The panel consists of six countries: Australia, Canada, France, Italy, the United Kingdom, and the United States. I have chosen these nations because of their status as advanced economies and global military powers. As determined by either development level (using the Human Development Index) or purchasing power parity GDP per capita, each of these countries is in the top thirty globally.[1] These countries are attractive choices in which to examine military, defense, and civil defense spending because of their ongoing commitments to international security and stability. Among G7 nations, Italy, France, the United King-

[1]HDI data are from the 2014 Human Development Index Report from the United Nations Development Program. PPP GDP data from World Bank, 2014.

dom and the United States are the top four nations as measured by the average of annual defense spending as a percentage of overall government spending over the period from 2000 to 2014.[2] Australia is included due to its participation in the Iraq War and, if it were in the G7, the third-highest defense spending metric just described. Canada is included due to its membership in the G7 and its participation, along with Australia, the U.K., and the U.S., in the 2001 invasion of Afghanistan. Together, the nations in the panel have a combined 240+ membership years on the United Nations Security Council and account for over $27 trillion in 2014 purchasing power parity GDP.

A typical entry into the collection of defense and civil defense news events consists of a description of the spending measure, the amount of spending, and the horizon of spending. As in Ramey (2011), the value of *defnews* for a quarter is the sum of present discounted values of all announcements during that quarter taken as a percentage of the previous quarter's nominal GDP. Ten-year nominal government bond yields are used to discount the announcement totals. As in Ramey (2011), announcements made during the last week of a quarter are recorded in the next quarter to more accurately reflect the time period in which the announcement affects the economy.

Many of the entries highlight changes scheduled to take place in the next or subsequent fiscal years, though some – generally those related to emergencies or geopolitical events at home or abroad – describe immediate increases or decreases in spending. In many cases, newspapers contain multiple accounts of these spending announcements. Wherever possible, I survey multiple reports for consistency and clarity of amounts and spending horizons. In very few cases across the entire news series do I include a spending announcement that has not been corroborated with more than one news report. I record only initial announcements of new, previously undisclosed spending episodes. In the event that the details of a spending plan are rumored or released prior to an official announcement, I record the earliest reliable account of the spending package.

[2] International defense spending data from SIPRI Military Expenditure Database 2015, http://milexdata.sipri.org.

Additionally, it is often the case that previously announced spending initiatives undergo revision, cancellation, upsizing, and downsizing due to a multitude of factors. In these cases, I record the present discounted value of the revision – positive or negative – in the period in which the revision was announced. As an example, in the second quarter of 2001, the government of Australia announced a plan to spend A$47 billion on new capital equipment projects for the Australian Defense Force over a ten-year period. In the first quarter of 2004, this spending plan was revised to add A$11.1 billion of new projects. In this case I recorded the present discounted value of the amount of the new projects as a positive shock in the first quarter of 2004. I do not retroactively edit the value of *defnews* for the second quarter of 2001, as this future change would have been unknown at the time of the original announcement.

The overwhelming majority of media reports for all nations outside of Australia are searched via Dow Jones' *Factiva* database.[3] The *Factiva* subscription I used provided access to more than thirty years of media coverage from thousands of news sources in more than twenty languages. Data from Australia was found using the *ProQuest Australia & New Zealand newsstand* database, a collection of more than thirty national and regional media sources. Although the database used for gathering data in Australia is different, each of the sources used to create values of the *defnews* variable for Australia is available in *Factiva*. I believe it is reasonable to assume that this idiosyncrasy in the data gathering process for Australia has no material effect on the results.

The process of querying these databases for the news reports needed to generate *defnews* involved developing a set of keywords or search terms, selecting a time frame in which to search, and selecting a collection of news sources to inspect. I describe this process in greater detail below.

The search terms used in conjunction with the databases mentioned above are listed in

[3]A small number (less than twenty) of news articles in the earliest years of the sample were found using the *ProQuest Newsstand* database rather than *Factiva*. Since the publication coverage of the two databases is very similar, it is unlikely this small inconsistency has any impact on the quality of the data or results.

Table 2.1. In Italy and France, the media reports we examined were printed primarily in terms of those countries' official languages. I translated the English versions of my keywords word-for-word into Italian and French using Google Translate. The Google Translate technology built into the Google Chrome web browser made reading these foreign language news articles quite easy. I occasionally came across phrases in the news articles that Google Translate was unable to translate back to English, as well as some idiomatic expressions whose English translation seemed illogical, but given my emphasis on corroborating across multiple sources, I do not believe this affects my results.

I searched my sample period in smaller intervals, often years, to break up the sample into pieces where it was easier to clue in on the flow of spending announcements and government actions in each nation. At several points in the sample, circumstances occurring in a particular country (e.g., a terrorist attack at home or in a neighboring country) concentrated the debate over spending proposals and military buildups in that nation around these high profile events. In these situations, I narrowed the search to a shorter period immediately following these high profile events to more carefully determine what actions, if any, were taken to change government spending levels.[4]

For each of the five countries in which I used Dow Jones' *Factiva* database to find news articles, each time a search was initiated I was required to select news sources to query. The publications selected are listed in Table 2.2. In each nation, I chose the two most widely circulated (as measured by daily volume) national broadsheet newspapers plus, depending on the nation, between one and three additional media sources. These additional sources may be other national newspapers (as in the case of the U.S., the U.K., and Italy), local newspapers (Canada), or newswire services (France). For Australia, the *ProQuest Australia & New Zealand newsstand* database searches its entire directory of media sources without requiring input from the user. For the U.S. *defnews* series, I use Ramey's data and augment it by adding civil defense-related spending announcements that occur during the sample

[4]An example of this would be France and the United Kingdom in the weeks following the July 7, 2005, London Underground bombings.

period.[5] No adjustments are made to Ramey's values of *defnews*, and I add only civil defense expenditures that are not already included in Ramey's data.[6]

I collected fifty-four quarters of *defnews* data for each nation in the panel. Of the 324 total country-quarter values, 119 are non-zero, meaning that over one-third of these country-quarters contain some announcement of defense and/or civil defense spending. Summary statistics for the *defnews* variable are given in Table 2.3. Graphs of the *defnews* time series for each nation are shown in Figure 2.1. All of the spending amounts, time frames, announcements, dates, and the sources used in constructing *defnews* are available in the data appendix.[7]

At several points in the data collection process I was forced to make judgment calls about the amounts or horizons for new spending plans, or regarding revisions to previously announced spending initiatives. The need to make these assumptions is the result of conflicting and incomplete news reports that are inconsistent in the ways they report expenditure announcements. It is often the case that the currency amounts appearing in headlines or the first few sentences of a news article contain amounts that include previously announced spending totals or aspects of spending programs that are not relevant to this analysis. Occasionally news articles will disclose large budget amounts without any further breakdown of what these figures include. Failure to further investigate these general accounts could dilute the accuracy of the *defnews* variable. I mitigate this by checking multiple news articles for more detailed accounts, which I often find, although in several situations I am left to make assumptions about spending totals in the absence of clear information.[8] When they exist, I also consult official sources and budget documents for further clarification. I test the sensitivity of the results to misinformation and errors in judgment by introducing measurement error into the values for *defnews*, similar to Ramey (2011). I find that it does not have a

[5] The Ramey data run through the end of 2013. I create 2014 data according to the same process as the other nations.

[6] Ramey's *defnews* series available at `http://econweb.ucsd.edu/~vramey/research/Ramey_News_US.xlsx`.

[7] Please contact the author for these details.

[8] All assumptions are disclosed in the data appendix. Please contact the author for details.

material impact on my findings.

2.4 Empirical Methodology

2.4.1 Model and Data

The goal of this empirical analysis is to determine the representative effects of changes in government spending on output, interest rates, consumption, unemployment, and the real wage. The data I have gathered are used in a panel VAR model similar to those discussed in Canova and Ciccarelli (2013). The nations in the panel are modern economies whose contributions to the global economy are significant. They are also nations in which we are likely to continue to see changes in government spending manifested in ways identical to what I have identified via *defnews*. To the extent that policymakers all over the world are concerned with global events in the context of near-term spending choices and economic forecasts, knowledge of the representative effects of expenditure changes reduces uncertainty in their decision-making process.

My model uses quarterly data from seven macroeconomic variables: the present discounted value of defense and civil defense news as a percentage of lagged nominal GDP (*defnews*), log of real government spending (*rgov*), log of real GDP (*gdp*), a three-month interbank rate adjusted for inflation (*tmo*), log of private final consumption expenditures (*consume*), unemployment rates for all persons age fifteen and older (*unempl*), and the log of an hourly index for manufacturing earnings (*wage*). All data other than *defnews* are collected from the OECD Main Economic Indicators database.[9] *defnews* is constructed according to the previous section. For each nation, the nominal GDP figures used to construct *defnews* are adjusted by the respective GDP deflator to generate quarterly real GDP.[10] Real government spending figures are generated by adjusting nominal government expenditures

[9] I obtained the data via the Federal Reserve Bank of St. Louis' FRED (Federal Reserve Economic Data) database.

[10] 2009 is the base year.

using the GDP deflator.[11] A nominal three-month interbank rate is adjusted using the GDP deflator annual rate of change to create a real interest rate. The other three variables are unadjusted from their original form other than taking logs.[12] These variables and their units are described in Table 2.4.

The panel VAR model pools observations from each nation. The resulting structure is similar to traditional VAR models that do not use panel data. Before aggregating across countries and time periods, the model is shown in (4.1). Column vector y_t contains current period (period t) values of $defnews, rgov, gdp, tmo, consume, unempl$, and $wage$ for nation i. Four quarters of lagged values of y_t are included.[13] A constant term a and dummy variables c_i for individual countries are also included. For each nation $i = 1, \ldots, N$, and for each time period $t = 1, \ldots, T$:

$$y_t' = a + y_{t-1}'\beta_1 + y_{t-2}'\beta_2 + y_{t-3}'\beta_3 + y_{t-4}'\beta_4 + c_i + u_t'. \tag{2.1}$$

To pool these data, I aggregate y_t', y_{t-j}', and u_t' over t into $\mathbf{y}_{t,i}$, $\mathbf{y}_{t-j,i}$, and $\mathbf{u}_{t,i}$, each of which contains all observations, across time, for a country i:

$$\mathbf{y}_{t,i} = \begin{bmatrix} y_{t,1}' \\ y_{t,2}' \\ \vdots \\ y_{t,T}' \end{bmatrix}, \quad \mathbf{y}_{t-j,i} = \begin{bmatrix} y_{t-j,1}' \\ y_{t-j,2}' \\ \vdots \\ y_{t-j,T}' \end{bmatrix}, \quad \text{and} \quad \mathbf{u}_{t,i} = \begin{bmatrix} u_1' \\ u_2' \\ \vdots \\ u_T' \end{bmatrix}.$$

I then combine each individual country's data into an aggregate sample. This step creates a representative economy. Each country's $\mathbf{y}_{t,i}$, $\mathbf{y}_{t-j,i}$, and $\mathbf{u}_{t,i}$ are stacked to form $\mathbf{Y}$, $\mathbf{X}$, and

[11] Nominal government expenditure data for Australia is not available. Instead, I use real government expenditure data and adjust it to be consistent with the other nations.

[12] The unemployment rate in France is calculated by dividing the population of unemployed persons age fifteen and older by the working age population ages fifteen to sixty-four. Only annual estimates of the latter quantity are available before 2003, so we assume that the quarterly figures are the same as the annual figures in these periods.

[13] AIC and BIC lag selection tests were done for models with lags of one, two, three, and four quarters. Calculations indicated that four lags are marginally better than one, two, or three lags.

$\mathbf{U}$. $\mathbf{X}$ includes the constant term and all lagged values of the variables in $\mathbf{Y}$.

$$\mathbf{Y} = \begin{bmatrix} \mathbf{y}_{t,1} \\ \mathbf{y}_{t,2} \\ \vdots \\ \mathbf{y}_{t,N} \end{bmatrix}, \quad \mathbf{X} = \begin{bmatrix} 1, \mathbf{y}_{t-1,1}, \ \ldots, \ \mathbf{y}_{t-4,1}, \\ 1, \mathbf{y}_{t-1,2}, \ \ldots, \ \mathbf{y}_{t-4,2}, \\ \vdots \\ 1, \mathbf{y}_{t-1,N}, \ \ldots, \ \mathbf{y}_{t-4,N}, \end{bmatrix}, \quad \text{and} \quad \mathbf{U} = \begin{bmatrix} \mathbf{u}_{t,1} \\ \mathbf{u}_{t,2} \\ \vdots \\ \mathbf{u}_{t,N} \end{bmatrix}.$$

The matrix $\mathbf{D}$ contains the country dummy variables to be used as additional regressors. Since the model contains a constant, $N-1$ country dummy variables are added. Formally:

$$\mathbf{D} = \begin{bmatrix} D_1 \\ \hline D_2 \end{bmatrix}$$

where $D_1 = I_{N-1} \otimes \iota_T$ has dimension $(N-1)T \times (N-1)$ and D_2 is a matrix of zeros with dimension $T \times (N-1)$. I_{N-1} is the $N-1$ square identity matrix and ι_T is a $T \times 1$ vector of ones. Combining the regressors such that

$$\tilde{\mathbf{X}} = \left[\ \mathbf{X} \,\middle|\, \mathbf{D} \ \right],$$

the global model can be expressed more compactly as:

$$\mathbf{Y} = \tilde{\mathbf{X}}\mathbf{B} + \mathbf{U}. \tag{2.2}$$

The matrix $\mathbf{B}$ contains coefficients for four lagged values of each variable, the constant term, and the coefficients on the dummy variables for the $N-1$ nations. Data from all six ($N = 6$) nations enter as both independent and dependent variables in (4.2). This is in contrast to other panel VAR studies (Love and Zicchino (2006) and Ciccarelli, Maddaloni, and Peydró (2013)) that divide the cross section into subgroups and estimate coefficients conditional on these subgroups. The sample period runs from the third quarter of 2001 to the fourth quarter of 2014. Also, the panel is balanced: accounting for lags, the dataset contains fifty ($T = 50$) quarters of data for each nation, generating an overall sample of 300 observations.

2.4.2 Prior Selection and Estimation

The classical method of estimating a panel VAR with fixed effects, pooled observations, and finite T is the generalized method of moments approach detailed in Arellano and Bond (1991). This technique requires differencing all variables in the model. Given the stacked nature of the data, differencing each variable for each nation in the panel would result in a loss of nearly ten percent of the sample. Since each country-quarter value of the one-of-a-kind *defnews* variable contains critical information about the the nature of unanticipated changes in defense spending, I turn to Bayesian methods to estimate the model without sacrificing data.

My seven-variable, four-lag VAR model contains over 200 parameters. Bayesian methods reduce the problem of finding point estimates for each element of $\mathbf{B}$ to selecting a small number of hyperparameters based the type of data in the system. The Bayesian principle updates the non-sample information contained in the prior with the explanatory power of *defnews* and my panel dataset to generate a posterior distribution that is used to analyze relationships between government spending and consumption, wages, and interest rates. The Normal-Wishart prior is used to estimate the model. This prior is a conjugate prior for the multivariate normal distribution, so both the prior and posterior distributions of the VAR coefficients are multivariate normal. The Normal-Wishart prior extends the Minnesota prior of Litterman (1986) by eliminating the assumption of a fixed diagonal variance-covariance matrix for the residuals u_t while maintaining computational simplicity. Direct sampling from the posterior is possible without use of Gibbs sampling or approximation techniques.

In describing the Normal-Wishart prior, I follow the notation used in Koop and Korobilis (2010).[14] The Normal-Wishart prior has the form

[14] I adapt the code found at `http://personal.strath.ac.uk/gary.koop/bayes_matlab_code_by_koop_and_korobilis.html` to estimate the model.

$$\alpha|\Sigma \sim \mathcal{N}(\underline{\alpha}, \Sigma \otimes \underline{V}) \tag{2.3}$$

where

$$\Sigma^{-1} \sim W(\underline{S}^{-1}, \underline{\nu}). \tag{2.4}$$

with $\mathcal{N}$ and W denoting the multivariate normal and Wishart distributions, respectively. In (4.3), $\alpha = vec(\mathbf{B})$ from the global model in (4.2). $\underline{\alpha}, \underline{V}, \underline{S}$, and $\underline{\nu}$ are hyperparameters whose values are chosen as follows. $\underline{\alpha}$ contains the prior means for the coefficients on the explanatory variables. Per Litterman (1986), I assign prior coefficient means based on the way each variable is measured. Variables measured in levels or log-levels are assumed to follow a random walk pattern. The initial values of the coefficients for each variable measured in levels or log-levels are one on the first own lag and zero elsewhere. For $defnews$, measured as a percentage of lagged nominal GDP, the prior mean on all coefficients is set to zero. The prior means of non-lagged variables (the constant and country dummies) are also set to zero.

To construct the prior variance-covariance matrix $\underline{V}$, I proceed according to Karlsson (2013) and choose a vector of three hyperparameters $\pi = (\pi_1, \pi_3, \pi_4)$.[15] π_1 controls overall prior tightness. As π_1 decreases in value, it increases the probability weight on the prior (relative to the likelihood function) in constructing the posterior distribution. π_3 controls the speed at which lagged values are shrunk to zero. As π_3 increases, the values of longer lagged variables have less probability weight in constructing the posterior distribution. π_4 controls the probability weight assigned to values of deterministic variables (the constant terms in A_0). I assign a very large probability weight to the information in the sample for the deterministic variables. Benchmark parameter settings for π are (.2, 2, 100,000). These settings are derived from initial values suggested in Canova (2007). I test several additional sets of hyperparameters that assign different levels of probability weight to the prior and sample information. These settings are described in Table 2.5. I do not modify the value

[15] π_2 controls the relative tightness on own lagged variables versus lags of other variables. In this implementation of the Normal-Wishart prior, this value is normalized to one.

of π_4 in any hyperparameter specifications given the lack of theoretical motivation in the literature for doing so.

$\underline{S}$ and $\underline{\nu}$ initialize the distribution from which the variance-covariance matrix Σ is drawn. The initial value for $\underline{S}$ is a diagonal matrix. Estimates of these diagonal entries come from univariate autoregressive models for each independent variable. I regress current period values of each independent variable on four quarters of own lagged values and a constant. The estimates of the error variance from each of these regressions form the entries of $\underline{S}$. $\underline{\nu}$ represents the degrees of freedom of the prior and is constructed according to Kadiyala and Karlsson (1997). In this model, $\underline{\nu} = 9$.

Applying Bayes' rule yields a multivariate normal posterior distribution from which I draw values of the coefficients in (4.2). The posterior has the form

$$\alpha|\Sigma \sim \mathcal{N}(\bar{\alpha}, \Sigma \otimes \bar{V}) \tag{2.5}$$

where

$$\Sigma^{-1}|y \sim W(\bar{S}^{-1}, \bar{\nu}). \tag{2.6}$$

$\bar{V}, \bar{\mathbf{B}}, \bar{S}$, and $\bar{\nu}$ are constructed analytically as follows:

$$\begin{aligned}
\bar{V} &= [\underline{V}^{-1} + \tilde{\mathbf{X}}'\tilde{\mathbf{X}}]^{-1} \\
\bar{\mathbf{B}} &= \bar{V}[\underline{V}^{-1}\underline{\mathbf{B}} + \tilde{\mathbf{X}}'\tilde{\mathbf{X}}\hat{\mathbf{B}}] \\
\bar{S} &= S + \underline{S} + \hat{\mathbf{B}}'\tilde{\mathbf{X}}'\tilde{\mathbf{X}}\hat{\mathbf{B}} + \underline{\mathbf{B}}'\underline{V}^{-1}\underline{\mathbf{B}} - \bar{\mathbf{B}}'\underline{V}^{-1} + (\tilde{\mathbf{X}}'\tilde{\mathbf{X}})\bar{\mathbf{B}} \\
\bar{\nu} &= NT + \underline{\nu}.
\end{aligned} \tag{2.7}$$

In (4.5), $\bar{\alpha} = vec(\bar{\mathbf{B}})$. In (4.7), $\underline{\mathbf{B}}$ is formed by transforming $\underline{\alpha}$ from a vector back into a matrix (i.e. back into the shape of $\mathbf{B}$), and $S = (\mathbf{Y} - \tilde{\mathbf{X}}\hat{\mathbf{B}})'(\mathbf{Y} - \tilde{\mathbf{X}}\hat{\mathbf{B}})$, where $\hat{\mathbf{B}}$ is the OLS estimate of $\mathbf{B}$.[16]

The structural shocks of the model are identified using a Choleski decomposition on Σ.

[16]Koop and Korobilis (2010), Kadiyala and Karlsson (1997), or Karlsson (2013) provide complete mathematical derivations of the items in (4.5), (4.6), (4.7).

I examine the responses of each variable and describe the results below.

2.5 Results and Interpretation

The impulse responses shown in Figures 2.2 to 2.8 are calculated by drawing coefficient values from the posterior distribution. On each set of axes I plot the median response from 10,000 posterior draws of α along with sixty-eight percent confidence bands.[17] Figure 2.2 shows the responses of each variable to a one percent shock to *defnews* using the benchmark hyperparameter settings. I find that a one percent increase in the present discounted value of defense and civil defense spending announcements leads to a significant increase in real government spending after approximately two quarters, with peak response approximately five quarters after the shock. This delay between the *defnews* shock and the observed increase in government spending speaks to the timing challenges described in Ramey (2011). Policy and implementation lags account for differences in the results of empirical studies identifying shocks with the traditional VAR approach versus military buildups and the narrative approach. The *defnews* variable captures the effects of government spending changes by pinpointing the initial mention of expenditure changes.

I find that the initial response of real GDP is negative. The real interest rate rises sharply on impact. Consumption decreases on impact and is significantly negative after approximately five quarters. Unemployment falls slightly on impact and increases steadily for six to eight quarters thereafter. Wages fall on impact but do not vary much over the five-year period after the *defnews* shock. The initial decline in GDP is likely due to the immediate decrease in consumption following an increase in government spending. For each of the six nations, consumption makes up the largest portion of output. Combined with the immediate decreases in wages and the unemployment rate, the results show modest evidence of the negative wealth effect predicted by the neoclassical model: upon hearing of future

[17] These confidence bands are constructed by plotting the 16th and 84th percentiles of the responses of each variable to *defnews* shocks.

increases in government spending, households reduce consumption and work more.

The gradual rise in the unemployment rate shown in these results is consistent with the neoclassical model's prediction that households supply more labor. Households supplying more labor means that some portion of these households is likely to be re-entering the labor force but unable to find immediate work. The sharp increase in interest rates may temper firms' demand for additional workers given the tightness in the loanable funds market that could be generated by increases in government spending financed by borrowing. Ideally, I would use data measuring hours worked to more firmly establish an increase in household labor supply, but these data are not available for all nations in the panel. This also has implications for wages, which do not respond meaningfully to the *defnews* shock. This could the result of wage stickiness or the fact that my measure of changes in labor market supply is somewhat of an approximation.

With the decline in the interest rate after its initial rise, the model may be capturing increases in saving by households. This spike in savings, which is consistent with the decline in consumption and GDP over the same period, may be the result of the uncertainty associated with the geopolitical events that generate many of the spending changes captured by *defnews*. As interest rates return to pre-shock levels, GDP and consumption begin to rise and unemployment falls.

The lack of significant responses throughout Figures 2.2 to 2.8 may be a product of the types of government spending identified via *defnews*. Defense and civil defense expenditures are more likely to be temporary relative to non-defense government spending for social programs or infrastructure. Indeed, Barro (1981) comments that military spending is temporary in nature. A significant portion of the spending done by Australia, Canada, the U.K., and the U.S. in the early years of my sample is related to these nations' role in the post-9/11 invasions of Afghanistan and Iraq. If the changes in defense and civil defense spending I have identified with *defnews* are temporary, this may explain the lack of significant impulse responses. Baxter and King (1993) find that the response of output to changes in government spending is smaller when the government spending changes are temporary. Aiyagari, Christiano, and Eichenbaum (1992) describe the response of output and unemployment to transitory movements in government expenditures as "small." Additionally, Aiyagari, Christiano, and Eichenbaum (1992) find that for both temporary and permanent changes to government spending, the response of the real interest rate is very small.

The negative movement in output following shocks to *defnews* could also be a consequence of the ways in which defense and civil defense news shocks are financed. Unfortunately there is not sufficient detail in the news articles used to construct *defnews* to determine whether new spending is financed by borrowing or taxes. In the latter case, the type of taxes levied may influence households' labor supply and consumption decisions. Baxter and King (1993) note that the lump-sum tax versus distortionary tax choice facing policymakers has substantial implications for empirical research. They find that output falls in response to government spending financed by income taxes since households are less willing to supply labor when the after-tax returns are lower. In several of the precedent empirical studies mentioned above, including Ramey (2011), marginal tax rates are included as a control for tax changes. Regrettably, data of this sort is not available for all of the nations in my panel.

Figures 2.3 to 2.7 show the impulse responses of the seven variables under different hyperparameter settings described in Table 2.5. The hyperparameters in these cases adjust overall prior tightness as well as the probability weight assigned to values of lagged variables. I find that the responses under each of these settings are nearly identical to those of the benchmark case. Figures 2.5 to 2.7 show impulse responses using an alternative version of $\underline{\alpha}$, the matrix of prior coefficient means. For this alternative version of $\underline{\alpha}$, those variables that have prior means of one on the first own lag also have prior means of .75, .5, and .25 on the second, third, and fourth own lags, respectively. This specification reflects a belief that these variables exhibit a stronger degree of persistence than the random walk model captures. I again find that the results are robust to these new hyperparameter settings.

Out of concern for errors in assumptions or other mistakes made during the data gathering process for defense and civil defense news, I test the sensitivity of my results to measurement error in *defnews*. I follow the procedure laid out in the appendix of Ramey (2011). This procedure adjusts the value of *defnews* in each quarter to account for mistiming by one quarter, past or future, of up to twenty percent of that quarter's value for *defnews*. This procedure simultaneously allows the value of *defnews* to be overestimated or underestimated by twenty percent. These responses are shown in Figure 2.8. They indicate that adding measurement error to *defnews* has little effect on the variable responses.

As an additional robustness check, I also test a smaller version of the model containing only five independent variables at a time. This reduction in the number of variables affects the nature of the multivariate normal posterior distribution from which the coefficients are drawn. In this specification I include the core variables *defnews*, *rgov*, *gdp*, and *tmo* with each of *consume*, *unempl*, and *wage* rotated in one at a time. Additionally, I test a model with all seven variables that also contains dummy variables marking the quarters spanning the global financial crisis. These dummy variables account for any idiosyncrasies particular to the periods from the fourth quarter of 2007 to the second quarter of 2009. All of these alternative specifications have no material effects on the results.

2.6 Concluding Remarks

Determining causality in the macroeconomy is a tall task for empirical research. In the absence of model economies or households willing to be the subjects of field experiments, the narrative approach is well-equipped to help establish some degree of cause and effect in the global macroeconomy. In the government spending literature, using the narrative approach can insulate researchers from questionable identifying assumptions of structural VAR analyses focusing on variation in total government spending. The Bayesian principle of combining a prior distribution with sample information is an intuitive method of weighing beliefs about the nature of macroeconomic data with a rich, unique dataset. I find moderate evidence in favor of the neoclassical growth model's prediction of a negative wealth effect that follows increases in government spending. Estimates of the response of output to government spending shocks indicate a value of the fiscal multiplier that is non-positive and close to zero. According to my results, policymakers eyeing fiscal policy as a tool to increase economic growth should carefully consider other options.

The degree of uncertainty that geopolitical events over the last fifteen years have injected into optimizing decisions of households may have significant consequences for the responses of macroeconomic variables to increases in government purchases. As Bloom (2009) has found, uncertainty shocks – which may be a side effect of the forces driving the changes in government spending identified via *defnews* – have substantial implications for employment and output. At any rate, the questions answered here are central to further empirical and theoretical study of advanced economies. Macroeconomists can safely assume that they will continue to be the subject of inquiry and exploration.

Country	**Keywords**
Australia	*australia AND (million OR billion) AND (defence OR military) AND (spending OR budget)*
Canada	*(million OR billion) AND (defence OR military OR security) AND (spending OR budget)*
France	*(securite OR terrorisme OR depenses) AND (defense OR militaire)*
Italy	*(milioni OR miliardi) AND (difesa OR militari OR sicurezza) AND (spesa OR bilancio)*
United Kingdom	*(million OR billion) AND (defence OR military OR security) AND (spending OR budget)*
United States	*(million OR billion) AND (defense OR military OR security) AND (spending OR budget)*

Table 2.1: Search terms used to construct the *defnews* variable

Country	**News Sources Searched**
Canada	*The Globe and Mail**, *National Post**, *Ottawa Citizen*, *Toronto Star*, *Vancouver Sun*
France	*Le Figaro**, *Le Monde**, *Agence French Presse*
Italy	*Corriere della Sera**, *La Repubblica**, *La Stampa*
United Kingdom	*The Daily Telegraph**, *The Times**, *Financial Times*, *The Guardian*, *The Independent*
United States	*The Wall Street Journal**, *The New York Times**, *Washington Post*, *USA Today*

Table 2.2: Media sources used in searches for spending announcements
Note: Starred sources are the top two most widely circulated national broadsheet newspapers in each respective country. Australia data is from *ProQuest Australia & New Zealand newsstand* database and requires no specifically named news sources to search.

Variable	**# Obs.**	**# Non-zero**	**# > |1%|**	**# > |2%|**	**Min**	**Max**
Australia	54	19	13	9	-5.8%	30.6%
Canada	54	22	9	6	-4.6%	4.7%
France	54	10	6	3	-0.3%	33.7%
Italy	54	21	1	1	-0.6%	2.3%
United Kingdom	54	27	8	2	-0.3%	6.3%
United States	54	20	16	12	-11.9%	20.3%

Table 2.3: Summary statistics for *defnews* variable

Variable	**Description**
defnews	PDV of defense/civil defense spending as a percent of lagged nominal GDP
gov	Log of real government expenditures
rgdp	Log of real GDP
tmo	Three-month interbank rates, adjusted for inflation
consume	Log of private final consumption expenditures
unempl	Unemployment rate of working population, age fifteen and up
wage	Log of manufacturing hourly wage index

Table 2.4: Variable definitions

Hyperparameter settings	**Description (relative to benchmark case π_b)**
$\pi_b = [.2, 2, 100, 000]$	Moderate tightness; inspired by Karlsson (2013) and Canova (2010)
$\pi_i = [.8, 2, 100, 000]$	More loose on coefficients; more weight given to sample data
$\pi_{ii} = [.05, 2, 100, 000]$	Tightest hyperparameter settings; most weight given to prior information
$\pi_{iii} = [.2, 1, 100, 000]$	Moderate tightness; assigns more weight to lagged values
$\pi_{iv} = [.8, 1, 100, 000]$	Loose prior; assigns more weight to lagged values
$\pi_v = [.05, 1, 100, 000]$	Tight prior; assigns more weight to lagged values

Table 2.5: Hyperparameter settings for model estimation

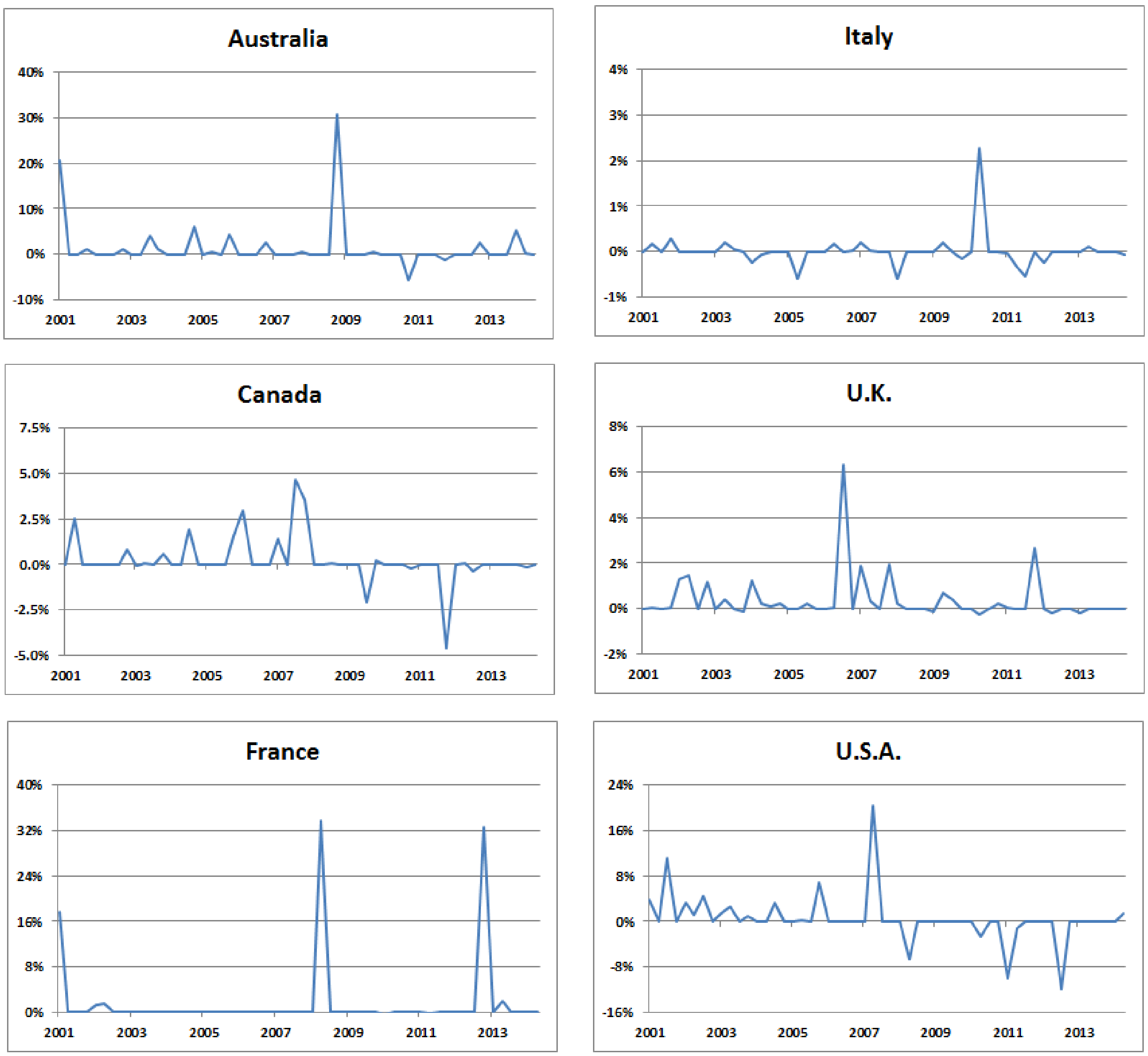


Figure 2.1: Plots of time series of *defnews*

Period is 3Q01 to 4Q14. *defnews* is measured as the present discounted value of the spending announcement as a percentage of the previous quarter's nominal GDP.

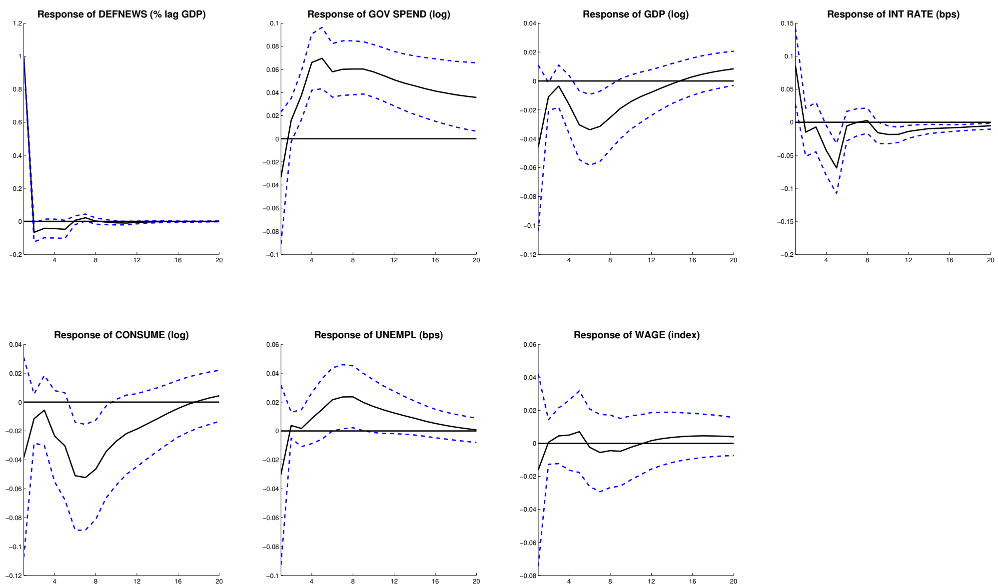


Figure 2.2: Responses to 1% shock to *defnews* with hyperparameter settings π_b
Prior coefficient means set to one on the first own lag and zeros elsewhere for all variables other than *defnews*, whose prior mean is zero everywhere. Sixty-eight percent confidence bands shown.

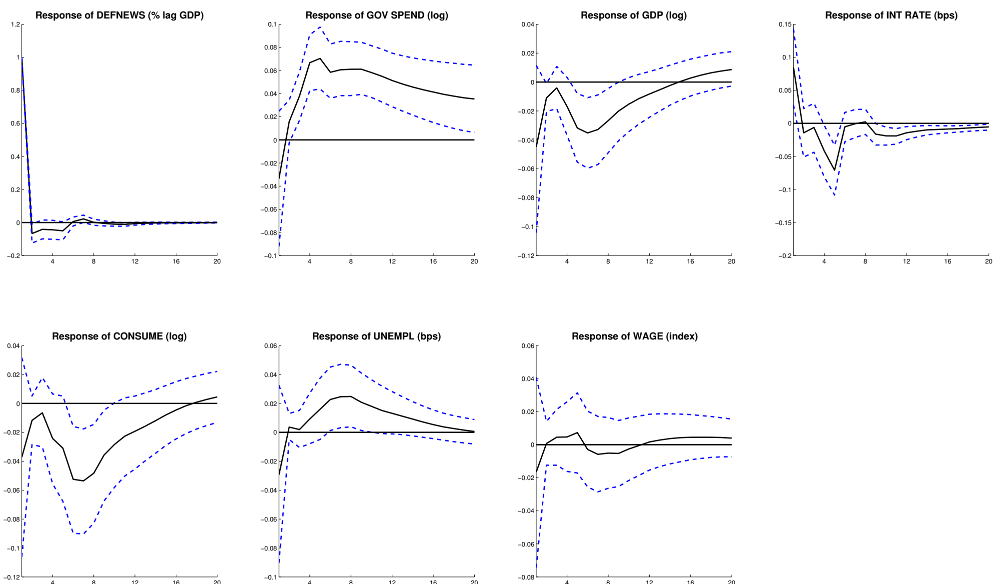


Figure 2.3: Responses to 1% shock to *defnews* with hyperparameter settings π_i
Prior coefficient means set to one on the first own lag and zeros elsewhere for all variables other than *defnews*, whose prior mean is zero everywhere. Sixty-eight percent confidence bands shown.

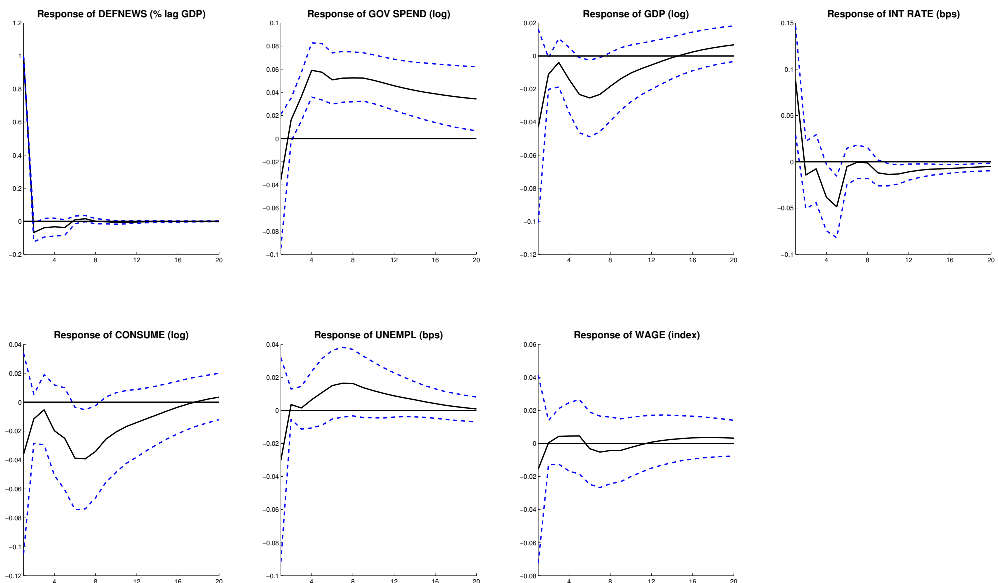


Figure 2.4: Responses to 1% shock to *defnews* with hyperparameter settings π_{ii}
Prior coefficient means set to one on the first own lag and zeros elsewhere for all variables other than *defnews*, whose prior mean is zero everywhere. Sixty-eight percent confidence bands shown.

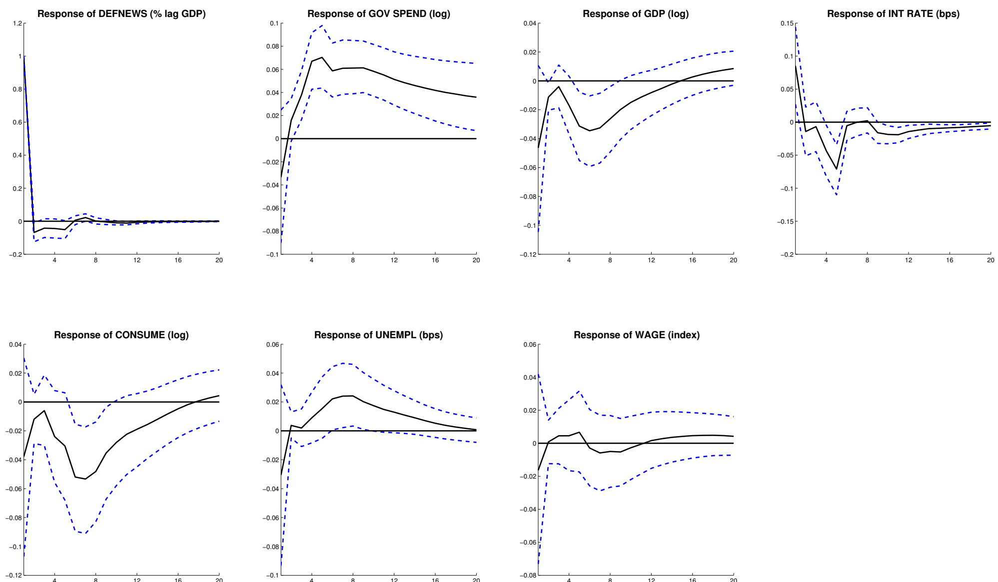


Figure 2.5: Responses to 1% shock to *defnews* with hyperparameter settings π_{iii}
Prior coefficient means set to 1, .75, .5, and .25 on the first, second, third, and fourth own lags (zeros elsewhere) for all variables other than *defnews*. Prior mean of *defnews* is zero everywhere. Sixty-eight percent confidence bands shown.

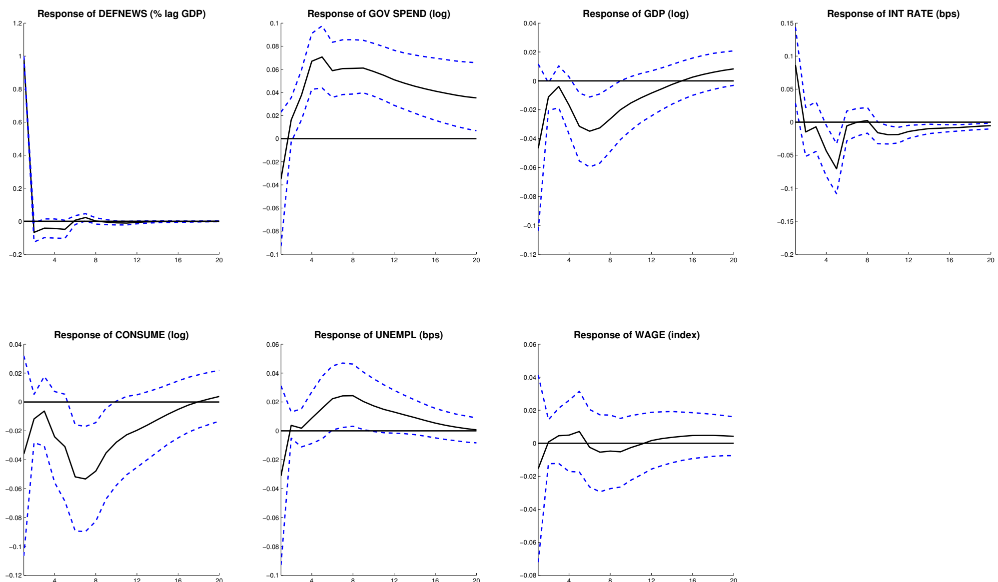


Figure 2.6: Responses to 1% shock to *defnews* with hyperparameter settings π_{iv}
Prior coefficient means set to 1, .75, .5, and .25 on the first, second, third, and fourth own lags (zeros elsewhere) for all variables other than *defnews*. Prior mean of *defnews* is zero everywhere. Sixty-eight percent confidence bands shown.

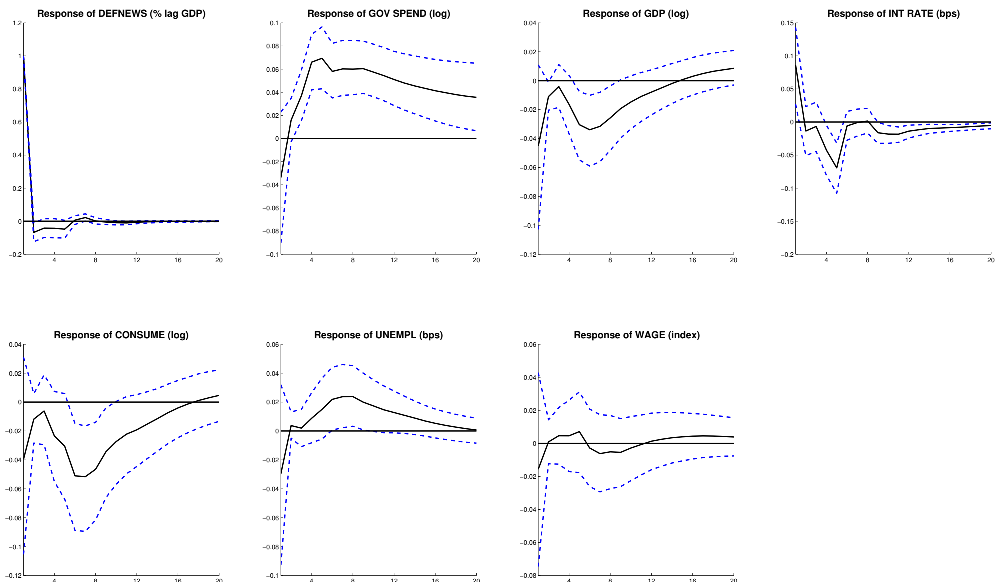


Figure 2.7: Responses to 1% shock to *defnews* with hyperparameter settings π_v
Prior coefficient means set to 1, .75, .5, and .25 on the first, second, third, and fourth own lags (zeros elsewhere) for all variables other than *defnews*. Prior mean of *defnews* is zero everywhere. Sixty-eight percent confidence bands shown.

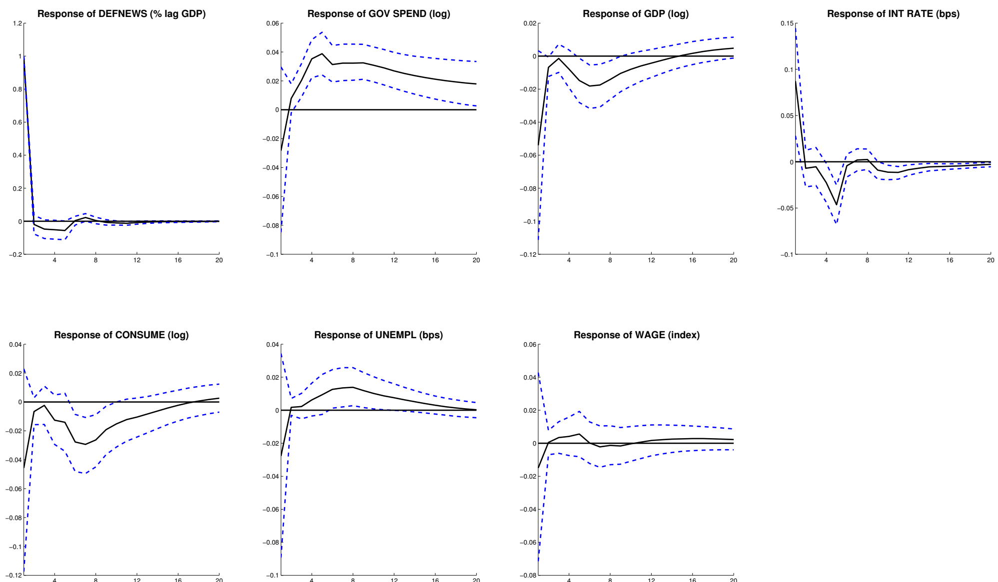


Figure 2.8: Responses to 1% shock to *defnews* with hyperparameter settings π_b *defnews* variable contains measurement error. Prior coefficient means set to one on the first lag and zeros elsewhere for all variables other than *defnews*, whose prior mean is zero everywhere. Sixty-eight percent confidence bands shown.

Chapter 3

Does Family Labor Supply Explain the Increase in Unemployment Duration?

3.1 Recent Developments in Labor Markets

A question central to economists' understanding of labor demand and labor supply is the relationship between the unemployment rate and unemployment duration. Headline measures of unemployment (i.e., the Bureau of Labor Statistics' "U3") are widely reported and scrutinized, but improvement in the aggregate unemployment rate can be achieved in spite of individuals whose unemployment spells drag on. Failure to account for these individuals may lead policymakers to overstate the health of the economy and turn their attention away from workers who face significant challenges in re-entering the labor market.

Abraham and Shimer (2002) document the evolving relationship between unemployment duration and unemployment rates in the United States. As unemployment rates rose in the 1970s, fluctuation in mean unemployment duration mirrored overall changes in unemployment rates into the early 1980s. Later in the decade, as the economy entered the 1990s, unemployment rates declined despite persistently high unemployment duration, each of which continued into the 2000s. The recessions of 2001 and 2007 to 2009 were marked by long recoveries and protracted unemployment durations. At the end of 2015, the headline unemployment rate had fallen to near five percent, in line with levels seen prior to the Great Recession. Mean unemployment duration, however, hovered around twenty-eight weeks – well above pre-2007 levels of seventeen weeks.

Concurrent with the the decoupling of unemployment rates and unemployment duration have been changes in family labor supply driven primarily by the entrance of women into the labor force. Blau and Kahn (2007) chronicle the rise in labor force participation of women over the period since the end of World War II until the year 2000. Labor supply of married women, along with women's real wages, increased significantly during the last two decades of the twentieth century. Women's labor force participation rose from thirty-one percent in 1947 to over sixty percent in 2001. Declining own-wage labor supply elasticities and continued efforts to close the gender pay gap are likely to have positive effects on the participation of women in the labor force.

We use data from the Annual Social and Economic Supplement (the March supplement) of the Current Population Survey (CPS) to examine the effects of spousal employment and family labor supply on unemployment duration and labor force status over the last three decades in the United States. We match data from individual CPS records to create measures of spouse wages, spousal employment, and family non-labor income. We use a proportional hazards model to calculate how these measures affect the length of unemployment spells and the likelihood of re-entering employment versus the likelihood of exiting the labor force. Despite our hypothesis that family labor income is used to subsidize long spells of unemployment, our results indicate that spousal employment does not decrease the hazard rates of exit from unemployment. This is the case for either exit to re-employment or exit from the labor force. These results hold after we separate the sample by gender and control for a host of demographic and socioeconomic characteristics, including unearned income and receipt of unemployment insurance.

We proceed as follows. Section 3.2 reviews several corners of the existing literature on unemployment duration and labor supply. Section 3.3 motivates this analysis using a collective labor supply model in which couples share income and individual labor supply decisions are influenced by the spouse's wage. Section 3.4 describes our data and empirical methodology in detail. Section 3.5 discusses our findings and Section 3.6 concludes.

3.2 Related Research from the Macro and Micro Literatures

This work contributes to both the macroeconomic and microeconomic literature by addressing unemployment duration, the added worker effect, and family labor supply. We describe the outstanding research on each topic below.

The dynamics of unemployment duration give policymakers information about the nature of unemployment. A decrease in headline unemployment rates in the face of unchanging or rising mean unemployment duration signals a substantial degree of long-term involuntary employment in the population rather than short-term frictional unemployment or movement in and out of the labor force. Changes in unemployment duration can have positive or negative effects on unemployed workers. Welfare is likely to fall if workers are risk-averse and increases in unemployment duration place additional uninsurable labor-income risk on workers. Long-term unemployment may place downward pressure on wages as chronically unemployed workers see their skills and job networks deteriorate. Workers may lower their reservation wage in hopes of ending their unemployment spell. On the other hand, if labor-income risks are self-insured through labor force participation by family or household members, this type of consumption smoothing could facilitate labor force attachment for long periods of time until workers find suitable job matches.

Abraham and Shimer (2002), Valletta (1998), and Valletta (2005) explore changes in the relationship between unemployment duration and unemployment rates since the late 1970s and find upward trends in unemployment duration. Abraham and Shimer attribute them to the a variety of factors. They find that a redesign of the Current Population Survey contributed by increasing mean unemployment duration by about half a week. Demographic changes in the U.S. population (primarily due to aging of baby boomers) place larger segments of the population into groups that experience longer unemployment spells. They also find that much of the overall rise in unemployment duration, particularly for women, is the result of increasing female labor force attachment and growing stability of female employment. Valletta (1998, 2005) attribute much of the recent rise in unemployment duration to increases in permanent job loss and struggles of labor force entrants in finding work. Valletta (2005) acknowledges the role of rising female labor force attachment, but notes that accelerated increases in unemployment duration during the second half of the 1990s – a period in which women's labor force behavior had converged to be similar to that of men (see also Blau and Kahn (2007)) – casts doubt on some of the findings of Abraham and Shimer.

As the global economy entered the Great Recession of the late 2000s, the average duration of unemployment spells in the United States reached all-time highs. By January 2010, mean unemployment duration was thirty weeks; in July 2011, a peak of forty weeks was reached. Aaronson, Mazumder, and Schechter (2010) estimate that about half of this run-up can be explained by the demographic changes highlighted by Abraham and Shimer. They also acknowledge extreme weakness in labor demand and, to a limited extent, expansion of unemployment insurance (UI) benefits at the state and federal levels as contributing factors. They show that lengthy unemployment spells are self-reinforcing: in a given period, individuals mired in long stretches of unemployment are less likely to find jobs in the next period. These would-be workers often remain unemployed long into the recovery.

The relationship between unemployment insurance and unemployment duration has been studied empirically by Moffitt (1985) and Meyer (1990). Each finds that the hazard rate of unemployment exit decreases when unemployment benefits rise. Chetty (2008) expands on these findings and describes how liquidity constraints affect unemployment search. Constrained workers who have greater difficulty smoothing consumption during unemployment spells reduce the urgency of their job search and use unemployment benefits to relax liquidity constraints. (Gruber (1997) finds that the unemployed make significant use of unemployment insurance for consumption smoothing.) Constrained workers are identified as those without working spouses and/or making mortgage payments. For unemployed workers that are not liquidity constrained, increases in unemployment duration after receiving benefits are related to the traditional moral hazard problem of reducing the cost of leisure. Overall, Chetty finds that approximately sixty percent of the marginal effects of unemployment benefits are related to liquidity concerns.

The severity of the economic downtown in 2007 to 2009 prompted expansion of unemployment insurance programs. Through the federal government's Extended Benefits and Emergency Unemployment Compensation programs, in the months during and after the Great Recession, workers were eligible for unemployment benefits for as long as ninety-nine weeks. UI eligibility requirements determined at the state level were expanded differently according to individual legislatures. Numerous papers have used these developments as motivation for new research. In several cases, researchers have used changes to UI to develop natural experiments centered around policy interventions. Farber and Valletta (2015) estimate the effects of benefit extensions on unemployment duration by creating a database of UI program dates and benefit levels across states. Information about maximum benefits, eligibility requirements, and dates of policy changes are used to generate indicator variables for extended benefits and periods in which an individual is in the last month of extended UI eligibility. They show that extended UI benefits reduce unemployment exits, a reflection of increased labor force attachment rather than reduced job finding. UI extensions are estimated to have small impacts on the aggregate labor market, increasing headline unemployment by less than half a percentage point. Howell and Azizoglu (2011) match this finding and posit that any increase in unemployment stemming from UI extensions is a reflection of workers remaining attached to the labor force for longer periods of time rather than enjoying higher incentives for leisure.

Hagedorn, Karahan, Manovskii, and Mitman (2013) focus on the responses of job and vacancy creation to expansion of unemployment benefits. In theory, expanded UI benefits lead workers to demand higher wages, shrinking firm profits and reducing the number of vacancies. They exploit state-by-state differences by comparing data from neighboring counties that belong to different states. Results indicate that UI extensions generate negative responses of vacancy creation and employment that are larger than the positive microeconomic effects related to worker attachment and job acceptance. The authors also find that counties with longer durations of unemployment benefits have higher unemployment levels.

Rothstein (2011) uses CPS data to identify the effects of UI extentions on unemployment duration during the Great Recession. To address endogeneity concerns about the impact of poor labor market conditions on the emergence of UI benefit extensions, Rothstein uses UI-ineligible workers as a control group and accounts for state-specific economic conditions to identify the effects of benefit extensions. Rothstein concludes that availability of extended UI benefits has small negative effects on hazard rates of exit from unemployment. The analysis suggests that both the headline unemployment rate and the long-term unemployment rate would have been lower by 0.2 and 1.6 percentage points, respectively, in the absence of extended unemployment benefits. Similar findings are reported in Fujita (2011) using monthly CPS data and in Nakajima (2011) using a structural model. Valletta and Kuang (2010) argue that extended UE benefits accounted for 0.4 of the more than five percentage point rise in unemployment during the Great Recession.

Hagedorn, Manovskii, and Mitman (2015) and Farber, Rothstein, and Valletta (2015) look at UI benefit rollbacks that took place in 2012 and 2013 as Congress ended the Emergency Unemployment Compensation program and states began to reduce UI eligibility. Farber, Rothstein, and Valletta (2015) estimate hazards of exit from unemployment in two periods with very different labor market conditions: 2008 to 2011 and 2012 to mid-2014. They find that extended UI benefits have significant negative effects on exits from employment to non-participation in the labor force. Hagedorn, Manovskii, and Mitman (2015) revisit their identification strategy of using data from adjacent counties sitting across state borders. They view the termination of extended UI benefits at the federal level as exogenous to state economic conditions and exploit discontinuities stemming from variation in state maximum UI benefit duration that remained after the federal policy change. They conclude that employment growth was higher in counties belonging to states where the decline in benefits was was larger than their neighbors. They suggest that the increases in employment are large enough to account for a majority of overall employment growth in the U.S. in 2014 and provide evidence that the response of employment to UI benefit reduction is positive.

This paper also adds to the literature describing the added worker effect. This term refers to the labor supply responses of married women to husband job loss (specifically, the propensity of wives to enter the labor force when their husbands become unemployed). This concept has been discussed at length in the labor economics literature, notably in Heckman and MaCurdy (1980) and Lundberg (1985). Lundberg (1985) finds a small positive response of female labor supply to husband job loss using household data from 1969 to 1973. Juhn and Potter (2007) use March CPS data from 1968 to 2005 to study labor market transitions of wives relative to their husbands' labor supply. They define the added worker effect as the difference between the probability of entering the labor force among wives whose husbands exit employment and the probability of entering the labor force among wives whose husbands remain employed. Their results suggest that the added worker effect remains relevant for some couples, but the that overall added worker effect has diminished since the 1990s due to sharp decreases in the number of wives not already employed. Mattingly and Smith (2010) find that the added worker effect was significantly larger during the Great Recession relative to the period leading up the the financial crisis of 2007 to 2009. Over the same period, Starr (2014) also finds significant increases in the probability of entering employment for wives with unemployed husbands. These probabilities were highest for younger wives and wives with non-adult children in the household. Starr finds evidence of a smaller added worker effect for husbands whose wives became unemployed during the Great Recession.

Stephens Jr. (2002) investigates the effects of husband job displacement (rather than unemployment) on wives' labor supply responses using Panel Study of Income Dynamics data. Stephens also looks at changes in spousal employment immediately before and after displacement occurs, and finds both pre- and post-displacement increases in wives' labor supply. Stephens finds that the nature of these responses depends on the magnitude of wages lost during displacement and the type of displacement (factory closings versus layoffs), suggesting that information received by couples in advance of and during displacement has implications for labor supply decisions. Cullen and Gruber (2000) look at the interaction between the added worker effect and unemployment insurance benefits and examine how spousal labor supply responds to receipt of non-labor income from unemployment insurance programs. They use Survey of Income and Program Participation (SIPP) data and find evidence of a substanding "crowding out" effect: if their husbands were not in receipt of UI benefits, wives' labor hours would be thirty percent higher. If spousal labor supply serves as a form of consumption smoothing for couples and families, these results suggest that there is some degree of substitutability among types of labor income insurance.

Our research also adds to the ongoing discussion about family labor supply and the role of families as economic decision-making units. Doepke and Tertilt (2016) discuss the importance of acknowledging how families make saving, investment, consumption, and labor supply decisions collectively. They argue that macroeconomic models focused on individuals may be inaccurately representing how agents make choices. Interactions among spouses, parents, and children, and potentially among families in the same household, have implications for macroeconomic research. Blundell and MaCurdy (1999) provide a comprehensive survey of family labor supply models and the literature up to the 1990s. Eggleston (2014) introduces a job search model in which married couples make consumption and labor supply decisions jointly. Mankart and Oikonomou (2015) develop a model where households are comprised of couples with primary and secondary earners. The model predicts that secondary earners participate in the labor market countercyclically, helping explain U.S. data that fail to show overall procyclical labor market participation. The behavior of secondary earners in this model helps two-member households insure against potential labor income losses. Guler, Guvenen, and Violante (2012) study how couples jointly search for work and how frictions unique to dual-earner households present new challenges relative to single-agent searches. Blundell, Pistaferri, and Saporta-Eksten (2016) document the importance of family labor supply as insurance against lost wages and find that over sixty percent of insurance against male wage loss comes from family labor supply.

3.3 Theoretical Motivation

To show how individual labor supply decisions could be influenced by spousal employment and wages, we turn to the collective labor supply model of Chiappori (1988, 1992). A two-adult household consisting of a married couple without children solves a family maximization problem. C_m and C_f denote male and female private consumption, respectively. L_m and L_f denote male and female leisure, respectively. The household's, or family's, utility is given by a function W taking individual male and female utilities U_m and U_f as arguments. The collective labor supply problem is to maximize a function that is separable in individual utilities subject to a household budget constraint:

$$\begin{aligned} &\max W\left[U_m(C_m, L_m), U_f(C_m, C_f)\right] \\ &\text{s.t. } C_m + C_f = M + W_m(T - L_m) + W_f(T - L_f) \end{aligned} \tag{3.1}$$

where M is non-labor income and T is total hours available for leisure or work. W_m and W_f denote male and female wages, respectively. Separability of the household utility function in individual consumption-leisure pairs permits consideration of the following individual utility maximization problems:

$$
\begin{aligned}
&\max U_m(C_m, L_m) \\
&\text{s.t. } C_m = M + W_m(T - L_m) - \phi(W_m, W_f, M)
\end{aligned}
\tag{3.2}
$$

and

$$
\begin{aligned}
&\max U_f(C_f, L_f) \\
&\text{s.t. } C_f = W_f(T - L_f) + \phi(W_m, W_f, M).
\end{aligned}
\tag{3.3}
$$

Chiappori (1992) shows that household decisions are Pareto efficient if and only if a sharing rule exists. The sharing rule ϕ appearing in each party's budget constraint illustrates how individual labor supply decisions depend on spousal wages in spite of individual utility functions that depend only on own consumption and leisure.

Using the job search model described in Rogerson, Shimer, and Wright (2005), spousal income sharing influences the reservation wage W_r. Individuals accept new job offers with wages $w \geq W_r$ and reject offers with $w < W_r$. The reservation wage W_r is given by

$$W_r = b + \frac{\alpha}{r} \int_{W_r}^{\infty} (w - W_r)\, dF(w) \tag{3.4}$$

which, after integrating by parts, is equal to

$$W_r = b + \frac{\alpha}{r} \int_{W_r}^{\infty} (1 - F(w))\, dw \tag{3.5}$$

where b is the benefit of not working, w is the wage associated with a new job offer, α is the arrival rate of new offers, and r is the interest rate. $1 - F(W_r)$ gives the probability of receiving an offer with $w \geq W_r$. As a result, multiplication by α gives the hazard rate of exit from unemployment $H = \alpha\,[(1 - F(W_r)]$. Since this hazard function is constant over time, the distribution of employment duration is exponential. Thus, we calculate the expected value of unemployment duration:

$$D = \int_0^\infty Hte^{-Ht}dt = \frac{1}{\alpha\left[(1-F(W_r)\right]} = \frac{1}{H}. \tag{3.6}$$

By rewriting the equation for the reservation wage given above in (3.5) as

$$\Phi(W_r, B, \alpha, r) = W_r - b + \frac{\alpha}{r}\int_{W_r}^\infty (1 - F(w))dw = 0, \tag{3.7}$$

we can take the derivative $\frac{\partial W_r}{\partial b} = -\frac{\Phi_b}{\Phi_{W_r}}$ and see that $\frac{\partial W_r}{\partial b} > 0$ and $\frac{\partial D}{\partial b} > 0$. This result confirms what is expected: as the benefit to not working b increases, so do the reservation wage and mean unemployment duration. In the collective model, the benefits to not working depend on the sharing rule. Ignoring the possibility of unemployment insurance or unemployment benefits, a married man who is not working with a spouse who is employed receives benefit $b_m = M - \phi(0, W_f, M)$. In a single-person household, or if his spouse were unemployed, this benefit would be simply $b = M$ such that $b_m > b$ (note that ϕ can be negative). We expect, therefore, that the average unemployment duration of a married man with an employed spouse is longer than if the man resided in a single person household.

3.4 Data and Empirical Methodology

Our data is from the Annual Social & Economic Supplement (known as the March supplement) of the Current Population Survey, a joint publication of the U.S. Census Bureau and the U.S. Bureau of Labor Statistics. The CPS collects labor supply information from surveys administered to a sample of approximately sixty thousand U.S. households. Each household is interviewed once a month for four months, and then once a month for four months in the same period next year. This survey structure – four months in the sample, eight months out, and four months in the sample again – facilitates month-over-month and year-over-year comparisons of weeks worked, hours per week, total income, and income components. The March supplement also includes work experience, employment status, migration, demographic, and occupation/industry information for each individual.

We use information contained in the 1979 to 2012 editions of the March supplement to study the length and determinants of unemployment spells. Since the labor market data published in one year actually pertain to the previous calendar year, our data are retrospective and correspond to the calendar years 1978 to 2011. The labor market metrics of primary interest to us are an individual's current labor force status (in or out, and employed or looking for work), weeks worked, weeks unemployed, and weeks not in the labor force over the previous year. Similar to Juhn, Murphy, and Topel (2002), we calculate time spent working as the number of weeks worked divided by fifty-two. Measures for time spent unemployed and not in the labor force are developed similarly.

Our measure of unemployment duration is equal to retrospective weeks unemployed for individuals who are currently employed. For individuals who are not working, unemployment duration is equal to the maximum of retrospective weeks unemployed and current weeks unemployed. Individuals may exit unemployment by finding a job or leaving the labor force. Those that remain unemployed at the end of the survey period are right-censored.

In addition to employment and labor force information, CPS survey respondents also report labor and non-labor income. We use these data to generate measures of weekly real wages.[1] Non-labor income and income from unemployment insurance programs are also measured. We link individual records in the CPS data to match spouses and create variables for spouse wages and spousal employment during the last twelve months. We also extract demographic information from the survey data for age, ethnicity, education level, home ownership, number of children, and gender. The sample is restricted to individuals ages twenty to sixty-four. Summary statistics are shown in Table 3.1.

We use the Cox proportional hazards model to quantify the effects of spouse wages and spousal employment on unemployment duration. This model allows us to estimate hazard rates of exit from unemployment given by

[1] Weekly wages are adjusted for inflation using the Consumer Price Index.

$$h(t) = h_0(t)\exp(\beta_1 x_1 + \cdots + \beta_k x_k) \tag{3.8}$$

where x_i are the covariates mentioned above and $h_0(t)$ is the baseline hazard. The two covariates of primary interest are spouse employed, a dummy variable equal to one if the individual's spouse was employed at any point during the last year, and spouse wage, which is constructed by matching individual records in the CPS data. Once again, our dependent variable, unemployment duration, is equal to the maximum of current weeks unemployed (active spells) and retrospective weeks unemployed (completed spells).

Models for exit to re-employment and exit from the labor force are estimated separately. Dummy variables for recessions in 1975, 1980, 1982, 1991, and 2001 are included in some specifications, along with interaction terms between these recession periods and dummy variables for spouse employment in the same period. We also use state fixed effects to control for differences in state unemployment insurance programs and the nature of state economies. Owing to papers in the literature by Farber and Valletta (2015) and Fujita (2011), our control for unemployment insurance is a binary variable that is equal to one if the individual received unemployment insurance in the last twelve months. Finally, we include a control for family non-labor income. This variable captures unearned income from a variety of sources, including veteran payments, worker's compensation, retirement, and investment income. We explain our findings in greater detail below.

3.5 Results and Interpretation

Table 3.2 shows estimates of the effects of spousal employment and various demographic and socioeconomic covariates on the hazard rate of exit to re-employment. In each specification of the model, the effect of spousal employment on the rate of exit to re-employment is either zero or positive. This result is inconsistent with our hypothesis that family labor supply insures against individual labor-income risk and allows unemployed workers to prolong their search for work. Estimates of the demographic and socioeconomic covariates, on the other hand, are rather intuitive: women and blacks are less likely to exit unemployment, whereas higher levels of education increase the probability of finding work. Individuals who own a home are also more likely to find work. More children, more nonlabor income, and receipt of unemployment insurance each decrease the hazard rate of exit from unemployment. The sign on the unemployment insurance dummy is consistent with a number of papers summarized in Section 3.2 that find negative effects of unemployment benefits on unemployment exits. For all covariates, the coefficients and the signficance of the coefficients are largely robust to model specification, including addition of dummy variables for recession periods, spousal employment during recession periods, and state fixed effects.

When we separate the sample by gender, the effects of spousal employment are much stronger for women. These results are shown in Table 3.3. The effects of spousal employment on exiting unemployment are insignificant for men but strongly significant and positive for women. Since women are more likely to be secondary earners, this result is surprising. The signs on the demographic covariates for both men and women are mostly similar to the full sample. For men, age decreases the likelihood of re-employment. For both sexes, receipt of unemployment insurance and family non-labor income increase unemployment duration. The effects of unemployment insurance are stronger for women.

Table 3.4 shows estimates of the effects of spousal employment and other covariates on the hazard rate of exit from the labor force. In the model with all covariates, we again see a counterintuitive result in which spousal employment increases the likelihood of exiting the labor force. If family labor supply allows unemployed workers to extend their job searches and stay connected to the labor force, we would expect the coefficient on spouse employed to be less than one. As was the case in Table 3.2, the coefficient on spouse employed either zero or positive for each specification. Older workers, black workers, and those with college degrees are less likely to leave the labor force. Women are much more likely than men to leave the labor force, presumably for family reasons. Individuals with more children are also more likely to leave the labor force. Receipt of unemployment insurance keeps unemployed workers attached to the labor force, but this is likely to be a response to UI eligibility requirements. Except for the variable measuring family non-labor income, we again find that the values and significance of the coefficients are robust to model specification and additional covariates.

Table 3.5 shows estimates of the hazard rate of labor force exit based on gender-specific subsamples. For both the men-only and women-only subsamples, the coefficients on spouse employed are not significantly different from zero. Black women are much less likely to exit the labor force than black men. For men, additional family non-labor income extends unemployment duration. As we would expect, the number of children has greater implications for women exiting the labor force than for men. For both sexes, older workers and workers with bachelor's degrees are more likely to stay connected to the labor force. The effects of unemployment insurance are similar for both men and women.

3.6 Conclusion and Future Research

Unemployment rates and unemployment duration are two key labor market metrics that command the attention of policymakers and economists. The dynamics of unemployment duration and the labor force have changed dramatically since the end of World War II. In the early 1950s, mean unemployment duration was around ten weeks and female labor force participation was less than thirty-five percent. Sixty years later, mean unemployment duration is upwards of twenty-five weeks, with over fifty percent of women in the labor force. In this paper we used CPS data and matched individuals to their spouses to determine how much of the upward trend in the length of unemployment spells can be explained by spousal employment. We motivated this question using a simple job search model that showed how individuals with employed spouses could be expected to stay unemployed for longer periods relative to single individuals and individuals in single-earner households.

After controlling for demographic factors, home ownership, unemployment insurance, and non-labor income, our findings do not support the claim that increases in spousal employment and the large entrance of women into the labor force in recent decades explain the decoupling of unemployment rates and unemployment duration. However, there are several areas where we can build on our findings in the near term. First, the period during and immediately after the Great Recession was characterized by significant changes in unemployment insurance and unemployment benefit programs. To the extent that these changes – namely, extension of UI eligibility up to as long as ninety-nine weeks – may have influenced the behavior of job seekers, this period warrants special consideration. Second, we plan to estimate our model using a competing risks framework that better manages the difference between unemployment exits for return to employment or exiting the labor force. Finally, some attention should be paid to the problem of selection bias and positive assortative mating in couples, an element that is gathering attention in the public debate over rising income inequality and has been linked to recent changes in female labor force participation by Greenwood, Guner, Kocharkov, and Santos (2014, 2016).

VARIABLES	N	Mean	SD	Min	Max
Spouse employed	137,266	0.812	0.391	0	1
Female	137,266	0.417	0.493	0	1
Age	137,266	38.9	11.0	20	64
Black	137,266	0.072	0.259	0	1
HS	137,266	0.398	0.490	0	1
Some college	137,266	0.215	0.411	0	1
Bachelor's	137,266	0.114	0.318	0	1
Grad	137,266	0.051	0.220	0	1
# of children	137,266	1.296	1.270	0	13
Own home	137,266	0.643	0.479	0	1
UI dummy	137,266	0.274	0.446	0	1
Family non-labor income	136,320	6,212	11,339	0	353,322

Table 3.1: Summary statistics for CPS data

VARIABLES	(1)	(2)	(3)	(4)	(5)
Spouse employed	1.147**	1.043**	1.044**	1.013	1.025*
	[0.00967]	[0.00924]	[0.0107]	[0.0104]	[0.0106]
Female		0.990	0.973**	0.982**	0.975**
		[0.00676]	[0.00668]	[0.00677]	[0.00672]
Age		0.996	1.003	0.997	1.001
		[0.00232]	[0.00235]	[0.00234]	[0.00235]
Age2		1.000**	1.000**	1.000**	1.000**
		[2.87e-05]	[2.88e-05]	[2.89e-05]	[2.90e-05]
Black		0.791**	0.766**	0.760**	0.762**
		[0.0103]	[0.0104]	[0.0103]	[0.0103]
HS		1.139**	1.167**	1.179**	1.183**
		[0.0103]	[0.0107]	[0.0108]	[0.0109]
Some college		1.244**	1.292**	1.305**	1.317**
		[0.0127]	[0.0134]	[0.0136]	[0.0137]
Bachelor's		1.366**	1.414**	1.448**	1.449**
		[0.0163]	[0.0171]	[0.0177]	[0.0177]
Grad		1.405**	1.421**	1.488**	1.470**
		[0.0220]	[0.0224]	[0.0238]	[0.0235]
# of children		0.967**	0.969**	0.970**	0.969**
		[0.00277]	[0.00278]	[0.00280]	[0.00280]
Own home		1.150**	1.168**	1.184**	1.187**
		[0.00850]	[0.00879]	[0.00898]	[0.00900]
UI dummy			0.782**		0.836**
			[0.00615]		[0.00715]
Family non-labor income				0.970**	0.978**
				[0.000932]	[0.00102]
Observations	137,266	137,266	137,266	136,320	136,320
Recession FE	NO	NO	YES	YES	YES
State FE	NO	NO	YES	YES	YES

seEform in brackets

** $p<0.01$, * $p<0.05$

Table 3.2: Estimated effects on probability of exit to re-employment, 1979 to 2012

VARIABLES	(1) ALL	(2) MEN	(3) WOMEN
Spouse employed	1.025*	1.015	1.140**
	[0.0106]	[0.0118]	[0.0276]
Age	1.001	0.993*	1.006
	[0.00235]	[0.00308]	[0.00372]
Age2	1.000**	1.000**	1.000**
	[2.90e-05]	[3.76e-05]	[4.66e-05]
Black	0.762**	0.756**	0.779**
	[0.0103]	[0.0134]	[0.0165]
HS	1.183**	1.179**	1.198**
	[0.0109]	[0.0134]	[0.0189]
Some college	1.317**	1.312**	1.334**
	[0.0137]	[0.0174]	[0.0231]
Bachelor's	1.449**	1.414**	1.486**
	[0.0177]	[0.0229]	[0.0286]
Grad	1.470**	1.461**	1.474**
	[0.0235]	[0.0310]	[0.0367]
# of children	0.969**	0.980**	0.955**
	[0.00280]	[0.00361]	[0.00447]
Own home	1.187**	1.206**	1.162**
	[0.00900]	[0.0118]	[0.0140]
UI dummy	0.836**	0.913**	0.742**
	[0.00715]	[0.0101]	[0.0102]
Family non-labor income	0.978**	0.971**	0.989**
	[0.00102]	[0.00131]	[0.00165]
Observations	136,320	79,558	56,762
Recession FE	YES	YES	YES
State FE	YES	YES	YES

seEform in brackets

** p<0.01, * p<0.05

Table 3.3: Estimated effects on probability of exit to re-employment, 1979 to 2012, by gender

VARIABLES	(1)	(2)	(3)	(4)	(5)
Spouse employed	1.370**	1.067*	1.057	1.022	1.065*
	[0.0340]	[0.0281]	[0.0320]	[0.0312]	[0.0326]
Female		3.009**	2.961**	2.999**	2.953**
		[0.0600]	[0.0593]	[0.0602]	[0.0593]
Age		0.849**	0.855**	0.849**	0.855**
		[0.00488]	[0.00495]	[0.00492]	[0.00496]
Age2		1.002**	1.002**	1.002**	1.002**
		[7.02e-05]	[7.06e-05]	[7.08e-05]	[7.09e-05]
Black		0.834**	0.824**	0.819**	0.823**
		[0.0285]	[0.0296]	[0.0295]	[0.0296]
HS		0.909**	0.937**	0.929**	0.937**
		[0.0214]	[0.0224]	[0.0223]	[0.0225]
Some college		0.976	1.029	1.004	1.028
		[0.0263]	[0.0282]	[0.0277]	[0.0284]
Bachelor's		0.765**	0.803**	0.794**	0.795**
		[0.0274]	[0.0291]	[0.0291]	[0.0291]
Grad		0.809**	0.819**	0.840**	0.810**
		[0.0397]	[0.0405]	[0.0420]	[0.0406]
# of children		1.060**	1.060**	1.062**	1.060**
		[0.00825]	[0.00825]	[0.00832]	[0.00828]
Own home		0.951*	0.974	0.969	0.970
		[0.0194]	[0.0202]	[0.0203]	[0.0203]
UI dummy			0.636**		0.630**
			[0.0149]		[0.0157]
Family non-labor income				0.983**	1.003
				[0.00267]	[0.00290]
Observations	137,266	137,266	137,266	136,320	136,320
Recession FE	NO	NO	YES	YES	YES
State FE	NO	NO	YES	YES	YES

seEform in brackets
** p<0.01, * p<0.05

Table 3.4: Estimated effects on probability of exiting the labor force, 1979 to 2012

VARIABLES	(1) ALL	(2) MEN	(3) WOMEN
Spouse employed	1.065*	0.973	1.029
	[0.0326]	[0.0402]	[0.0496]
Age	0.855**	0.880**	0.872**
	[0.00496]	[0.00922]	[0.00644]
Age2	1.002**	1.002**	1.002**
	[7.09e-05]	[0.000122]	[9.33e-05]
Black	0.823**	1.012	0.733**
	[0.0296]	[0.0600]	[0.0333]
HS	0.937**	0.888**	0.958
	[0.0225]	[0.0354]	[0.0290]
Some college	1.028	1.079	0.990
	[0.0284]	[0.0495]	[0.0344]
Bachelor's	0.795**	0.732**	0.802**
	[0.0291]	[0.0480]	[0.0358]
Grad	0.810**	0.840*	0.775**
	[0.0406]	[0.0694]	[0.0489]
# of children	1.060**	1.009	1.077**
	[0.0828]	[0.0138]	[0.0103]
Own home	0.970	0.926*	0.972
	[0.0203]	[0.0339]	[0.0247]
UI dummy	0.630**	0.655**	0.615**
	[0.0157]	[0.0264]	[0.0196]
Family non-labor income	1.003	1.015**	1.001
	[0.00290]	[0.00516]	[0.00356]
Observations	136,320	79,558	56,762
Recession FE	YES	YES	YES
State FE	YES	YES	YES

seEform in brackets
** p<0.01, * p<0.05

Table 3.5: Estimated effects on probability of exiting the labor force, 1979 to 2012, by gender

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