Essays on Business Cycles and Stabilization Policy

by

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To my parents and to Nitya.

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ABSTRACT

Essays on Business Cycles and Stabilization Policy

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Chair: Christopher L. House

This dissertation is a collection of essays on fiscal policy, monetary policy and the international transmission of business cycle shocks. Chapter 1 highlights the importance of distinguishing durable and nondurable goods when conducting countercyclical fiscal policy. It shows that the fiscal multiplier for purchases of durable goods is much smaller than the multiplier for nondurable goods. Standard models predict small durables multipliers because private sector purchases of durable goods are highly intertemporally substitutable and therefore easily crowded out. Empirical estimates based on U.S. data confirm this result. In aggregate time series data output rises by about 50 cents less if the government purchases 1$ of durable rather than nondurable goods. Industry-level estimates also point to smaller durable goods multipliers. The findings of this chapter suggest that infrastructure spending which is frequently part of fiscal stimulus packages is relatively ineffective at raising aggregate demand.

Chapter 2, joint with Christopher L. House, shifts focus to monetary policy and analyzes the optimal Taylor rule in a standard New Keynesian model. If the central bank can observe

the output gap and inflation without error, then it is typically optimal to respond infinitely strongly to observed deviations from the central banks targets. If it observes inflation and the output gap with error, the central bank will temper its responses so as not to impart unnecessary volatility to the economy. If the Taylor rule is expressed in terms of estimated output and inflation then it is optimal to respond infinitely strongly to estimated deviations from the targets. Under such a Taylor rule the estimates of inflation and the output gap should be perfectly negatively correlated. In the data, inflation and the output gap are weakly correlated, suggesting that the central bank is systematically underreacting.

Chapter 3, joint with Aaron Flaaen and Nitya Pandalai-Nayar, studies the cross-country transmission of shocks. Using the 2011 Tōhoku earthquake as a natural experiment, the study shows that firms reliant on Japanese intermediates experienced significant drops in production after the disruption. These findings imply that supply chains are sufficiently inflexible to play an important role in the international transmission of shocks.

CHAPTER I

Government Spending and Durable Goods

1.1 Introduction

When the government raises spending by one dollar, GDP changes by an amount that is generally *not* one dollar. The difference arises from the fact that households and firms change their behavior when the government intervenes—a phenomenon known as crowding in or out. It is well understood that the private sector response depends on whether the fiscal intervention is temporary or permanent (*Baxter and King* 1993). Recent research has suggested that the response can also depend on whether the economy's factors of production are underutilized (e.g. *Michaillat* 2014) and whether the central bank accommodates the change in spending (e.g. *Christiano et al.* 2011, *Woodford* 2011, and *Rendahl* 2014).

In this paper I argue that the composition of government spending matters for the size of the fiscal multiplier. I show that in a two sector model purchases of durable goods have a much smaller multiplier than purchases of nondurable goods. I then estimate multipliers separately for spending in durables and nondurables industries. While, on average, a dollar of spending on durable goods raises industry gross output by less than 30 cents, gross output rises more than one-for-one if the government buys nondurable goods. These findings suggest that infrastructure spending has very small affects on aggregate demand.

The demand for nondurable goods differs fundamentally from the demand for new durables. Whereas nondurable goods and services are immediately consumed, durable goods such as cars, appliances or structures have service lives of many years. If the price of a durable good rises temporarily, households can rely on the existing stock and postpone new purchases until prices revert to lower levels. All else equal, a longer service life leads to a greater intertemporal substitutability of purchases.

Intertemporal substitutability is key for the fiscal multiplier because it determines how much private sector spending is crowded out. If the government temporarily raises spending and drives up prices, households delay new purchases until the fiscal expansion ends. The higher the intertemporal elasticity of substitution the larger the degree of crowding out. It follows that multipliers for durable goods with high intertemporal elasticities of substitution are smaller than multipliers for nondurable goods with low intertemporal elasticities of substitution.

In standard models the difference between durable and nondurable multipliers is large. When the service life is calibrated to realistic levels, the durables multiplier is often less than one third of the nondurables multiplier. I also show that the durables multiplier approaches zero as the service life of the durable good becomes large. The prediction that durables multipliers are smaller than nondurables multipliers holds for both New Keynesian and neoclassical models.

Research by *Woodford* (2011), *Christiano et al.* (2011), and *Farhi and Werning* (2012), among others, has emphasized that the size of the fiscal multiplier is highly sensitive to the monetary policy response. Interestingly, the durables multiplier often depends less on monetary policy than the nondurables multiplier. Because purchases of durables crowd out private sector spending, equilibrium quantities and prices move little relative to the case in which the government purchases nondurable goods. With output and prices barely changed, the monetary policy response matters less. In the special case in which the service life of durables becomes large, the allocation is independent of monetary policy for a large class of rules.

I estimate fiscal multipliers separately for durable and nondurable goods, beginning with evidence based on U.S. national accounts data. In simple regressions which take U.S. military spending as exogenous (*Hall* 2009 and *Barro and Redlick* 2011) the nondurables multiplier exceeds the durables multiplier by about 0.5. Consistent with theory, government purchases of durable goods are associated with declines in private sector investment and purchases of durable goods.

I subsequently turn to an analysis at the industry level. In particular, I assemble a new dataset from the NBER-CES Manufacturing Industry Database and two databases covering the universe of U.S. military prime contracts. The merged dataset contains industry-level outcomes such as gross output, value added, and employment together with military spending on goods in each 4-digit SIC manufacturing industry.

As predicted by the theory, industries that produce durable goods respond much less to fiscal expansions than industries that produce nondurable goods. Using an identification strategy closely related to *Ramey and Shapiro* (1998), *Nekarda and Ramey* (2011), and *Nakamura and Steinsson* (2014), I first estimate impulse response functions. I then construct industry-level multipliers from the estimated impulse response functions separately for durable and nondurable goods. The multipliers for all five measures of economic activity I consider—gross output, value added, cost of materials, energy expenditures, and employment—are smaller if spending takes place in durable goods industries.

This paper is related to a large literature on the effects of government spending. *Hall* (2009) shows in a simple static model with a single nondurable good that the fiscal multiplier is decreasing in the intertemporal elasticity of substitution. In his model the multiplier tends to zero as the intertemporal elasticity of substitution approaches infinity. I show that the same result holds approximately for short-run multipliers of long-lived durable goods.

Consistent with this result, *House* (2009) argues that among the options that households face in response to greater government spending, reducing investment is more attractive than reducing consumption or raising the supply of labor.

Most of the literature on fiscal multipliers does not distinguish between purchases of durable and nondurable goods. Exceptions include *Perotti* (2004) and *Pappa* (2009a,b) who study the effects of government consumption and investment using structural vector autoregressions. Their findings are broadly consistent with the hypothesis that government consumption is associated with greater multipliers than government investment.[1] Paying particular attention to anticipation effects and timing *Leduc and Wilson* (2013) study the effects of public infrastructure spending. They find a very large short-term multiplier in recessions and an insignificant short-term multiplier in expansions.[2]

A number of studies have suggested that the fiscal multiplier is larger during economic downturns than in times of peak economic activity (e.g. *Auerbach and Gorodnichenko* 2012, 2013, *Bachmann and Sims* 2012, and *Michaillat* 2014).[3] By far the most commonly cited reason for such state dependence is that the presence of slack: If factors of production are underutilized there is less crowding out of private sector spending, leading to a larger fiscal multiplier. Since durable goods industries exhibit high volatility over the business cycle, one may conjecture that the state dependence hypothesis applies particularly to spending on durable goods. Yet, when I allow for state dependence in the estimation of the durable goods multiplier, I find no indication for greater effectiveness when spending occurs in times of economic slack.[4]

[1] See in particular tables 4 to 6 in *Perotti* (2004), the "typical" state-level employment responses in figure 3 in *Pappa* (2009a) and table 1 in *Pappa* (2009b).

[2] Other papers that distinguish different government interventions include *Finn* (1998), and *Pereira* (2000). *Aschauer* (1989) studies the effect of public expenditure on productivity.

[3] The evidence on state dependence of fiscal multipliers is not uncontroversial. *Owyang et al.* (2013) find that the multiplier in recessions is large in Canada, but not in the U.S.

[4] In fact, *Berger and Vavra* (2014) argue that the durable goods fiscal multiplier may be *smaller* in recessions than during expansions.

Although I cast my analysis largely in terms of consumer durables, I note that the demand for investment goods such as machines, ships or structures is also highly elastic (*House and Shapiro* 2008). In the empirical part of this paper, I will refer to all long-lived goods as durable goods, regardless of whether they are consumer durables or investment goods. Because services are immediately consumed, they are best understood as nondurable goods in the context of this paper.

The remainder of the paper is structured as follows. In the next section, I lay out a simple theory predicting that the multiplier for durables is small relative to that of nondurable goods. I then turn to the empirical analysis in Section 1.3. In Section 1.4 I analyze an extension of the model in which the interest rate is pinned at the zero lower bound. Section 1.5 concludes.

1.2 Theoretical analysis

I next present a New Keynesian two sector model to study the effectiveness of government spending on durable and nondurable goods. After describing the model, I demonstrate that the intertemporal elasticity of substitution of durable goods purchases is much larger than that of nondurable goods purchases. I then show that their high intertemporal substitutability renders the durable goods multiplier small.

1.2.1 Model description

The model is designed to mirror the empirical setting I face below: There is a large number of (4-digit SIC) industries. Some of these industries produce durable goods, others produce nondurable goods. I am interested in the effect on GDP when the government increases spending in a typical durable or a typical nondurable goods industry.

To keep the framework tractable, I only model two sectors, a small sector X and a large sector Z. The small sector represents a typical 4-digit SIC manufacturing industry that is subjected to a government spending shock. Depending on the choice of the depreciation rate δ_D, the small sector produces either durable or nondurable goods for final consumption. The large sector Z represents the aggregate of all remaining industries. It produces goods which can alternately be used for final nondurable consumption C, for investment into capital of the two sectors, I_Z and I_X, or for intermediate goods M_X used in the production of good X. For simplicity, there is no government spending on goods in the Z sector. Notice that Z is a hybrid sector, producing both nondurable goods (C and M_X) and durable investment goods (I_Z and I_X).

1.2.1.1 Representative household

The representative household maximizes life-time utility

$$\mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \left[u\left(C_t, D_{H,t}\right) + \Gamma_t - v\left(N_{X,t}, N_{Z,t}\right) \right]$$

subject to the nominal budget constraint

$$\begin{aligned} &P_{X,t} X_{H,t} + P_{Z,t}\left(C_t + I_{X,t} + I_{Z,t}\right) + B_t + T_t \\ &\qquad = W_{X,t} N_{X,t} + W_{Z,t} N_{Z,t} + R_{X,t} K_{X,t-1} + R_{Z,t} K_{Z,t-1} + B_{t-1}\left(1 + i_{t-1}\right) + \Pi_t, \quad (1.1) \end{aligned}$$

the accumulation equations

$$D_{H,t} = X_{H,t} + \left(1 - \delta_D\right) D_{H,t-1}, \quad K_{j,t} = I_{j,t} + \left(1 - \delta_K\right) K_{j,t-1}, \; j \in \{Z, X\}, \qquad (1.2)$$

and a no-Ponzi game condition.

Utility is derived from three components. The first component, u, reflects the benefit derived from the consumption of the nondurable good C_t and the good $D_{H,t}$. When $\delta_D = 1$, $D_{H,t} = X_{H,t}$ is a nondurable good. For $\delta_D < 1$, $D_{H,t}$ is the household's stock of durable goods, while $X_{H,t}$ are new purchases. The second component, Γ_t, represents the household's utility derived from government purchases. I assume that Γ enters additively separable so that government spending is neither a substitute nor a complement for private sector spending. Finally, the third term is the household's disutility from supplying labor to the two sectors. Letting subscripts denote partial derivatives, I assume that 1) $u_C, u_D, v_X, v_Z > 0$, 2) $u_{CC}, u_{DD} < 0$, 3) $v_{XX}, v_{ZZ} > 0$, and 4) that Inada-type conditions hold.

The remaining notation is chosen as follows. $P_{X,t}$ and $P_{Z,t}$ denote the prices in the X and the Z sector. Wages are analogously denoted by $W_{X,t}$ and $W_{Z,t}$. Each sector has its own capital stock $K_{X,t}$ and $K_{Z,t}$, earning rental rates $R_{X,t}$ and $R_{Z,t}$. The representative household can hold risk-free nominal bonds, B_t, paying interest rate i_t. Π_t are profits and T_t is a lump-sum tax.

1.2.1.2 Firms

Both sectors have a representative aggregating firm and a unit continuum of differentiated firms. The aggregating firms assemble the differentiated varieties into CES bundles

$$X_t = \left[\int_0^1 x_t(s)^{\frac{\varepsilon-1}{\varepsilon}} \, ds \right]^{\frac{\varepsilon}{\varepsilon-1}}, \quad Z_t = \left[\int_0^1 z_t(s)^{\frac{\varepsilon-1}{\varepsilon}} \, ds \right]^{\frac{\varepsilon}{\varepsilon-1}}. \tag{1.3}$$

Optimal behavior in competitive markets implies the demand functions

$$x_t(s) = X_t \left(\frac{p_{x,t}(s)}{P_{X,t}} \right)^{-\varepsilon}, \quad z_t(s) = Z_t \left(\frac{p_{z,t}(s)}{P_{Z,t}} \right)^{-\varepsilon} \tag{1.4}$$

where $p_{x,t}(s)$ and $p_{z,t}(s)$ denote the prices of a generic variety s in each sector and $P_{X,t}$ and $P_{Z,t}$ are given by

$$P_{X,t} = \left(\int_0^1 \left(p_{x,t}(s)\right)^{1-\varepsilon} ds \right)^{\frac{1}{1-\varepsilon}}, \quad P_{Z,t} = \left(\int_0^1 \left(p_{z,t}(s)\right)^{1-\varepsilon} ds \right)^{\frac{1}{1-\varepsilon}}. \tag{1.5}$$

A differentiated firm in sector X produces variety s using production function

$$x_t(s) = \left[\left(k_{x,t}(s)\right)^{\alpha} \left(n_{x,t}(s)\right)^{1-\alpha} \right]^{\chi} \left[m_{x,t}(s)\right]^{1-\chi}. \tag{1.6}$$

The firm rents capital $k_{x,t}(s)$ at rate $R_{X,t}$ and employs labor $n_{x,t}(s)$ at wage $W_{X,t}$. Additionally, production requires an intermediate $m_{x,t}(s)$ from the large sector. Parameter χ is the cost share of capital and labor and $(1-\chi)$ is that of intermediates. Cost minimization in competitive factor markets yields the firm's conditional factor demand functions and an expression for its marginal costs $MC_{X,t}$.

Firms set prices as in *Calvo* (1983). Let θ_X denote the probability that a firm in sector X cannot adjust its price. Let further λ denote the Lagrange multiplier on the budget constraint (1.1). Then the monopolistically competitive firm chooses the reset price $p^*_{X,t}$ to maximize objective

$$\mathbb{E}_t \sum_{j=0}^{\infty} (\theta_X \beta)^j \frac{\lambda_{t+j}}{\lambda_t} \left[p^*_{X,t} x_{t+j} - MC_{X,t+j} x_{t+j} \right]$$

subject to the sequence of demand functions (the first equation in 1.4) and its marginal costs $MC_{X,t+j}$. The optimal reset price is

$$p^*_{X,t} = \frac{\varepsilon}{\varepsilon - 1} \frac{\mathbb{E}_t \sum_{j=0}^{\infty} (\theta_X \beta)^j \lambda_{t+j} X_{t+j} \left(P_{X,t+j}\right)^{\varepsilon} MC_{X,t+j}}{\mathbb{E}_t \sum_{j=0}^{\infty} (\theta_X \beta)^j \lambda_{t+j} X_{t+j} \left(P_{X,t+j}\right)^{\varepsilon}}.$$

Monopolistic competitors in the large sector behave similarly. The only difference is that they do not require an intermediate input. Their production function is simply $z_t(s) = (k_{z,t}(s))^{\alpha}(n_{z,t}(s))^{1-\alpha}$, and capital and labor are paid the rental rate $R_{Z,t}$ and the wage $W_{Z,t}$. Price rigidity in sector Z is parameterized by θ_Z. Notice that I assume that productivity in both sectors is unaffected by government spending.

1.2.1.3 Market clearing, government, accounting, and monetary policy

Market clearing in sectors Z and X requires

$$Z_t = C_t + I_{X,t} + I_{Z,t} + M_{X,t} \text{ and } X_t = X_{H,t} + X_{G,t} \tag{1.7}$$

where $M_{X,t} = \int_0^1 m_{x,t}(s)\,ds$ is the total of intermediates demanded by the X sector and X_G is government spending on good X. The labor and capital market clearing conditions are given by

$$N_{X,t} = \int_0^1 n_{x,t}(s)\,ds, \quad N_{Z,t} = \int_0^1 n_{z,t}(s)\,ds \tag{1.8}$$

and

$$K_{X,t-1} = \int_0^1 k_{x,t}(s)\,ds, \quad K_{Z,t-1} = \int_0^1 k_{z,t}(s)\,ds. \tag{1.9}$$

Since the economy is closed and the government always balances its budget ($T_t = P_{X,t}X_{G,t}$), bonds are in zero net supply, $B_t = 0$.[5] I assume that government purchases in sector X follow the AR(1) process

$$X_{G,t} = (1-\varrho_X)\,X_G + \varrho_X X_{G,t-1} + \varepsilon_{G,t}. \tag{1.10}$$

Variables without time subscripts, such as X_G, denote steady state values.

[5]Since Ricardian equivalence holds in this model, the balanced budget assumption is not restrictive.

GDP in constant prices is

$$Y_t = P_X X_t - P_Z M_{X,t} + P_Z Z_t. \tag{1.11}$$

The GDP deflator and inflation are then

$$P_t = \frac{P_{X,t} X_t + P_{Z,t} Z_t - P_{Z,t} M_{X,t}}{P_X X_t + P_Z Z_t - P_Z M_{X,t}} \text{ and } \pi_t = \left(P_t - P_{t-1}\right)/P_{t-1}. \tag{1.12}$$

I initially assume that the monetary authority follows a fairly general rule of the form

$$i_t = \iota\left(\left\{i_{t-s-1}, P_{t-s}, P_{X,t-s}, P_{Z,t-s}, Y_{t-s}, X_{t-s}, Z_{t-s}\right\}_{s=0}^{\infty}\right). \tag{1.13}$$

where ι is any function of the given arguments. This completes the description of the model. I summarize all equations in Appendix A.1.1.

1.2.2 The demand for durable and nondurable goods

I next compare the demand for durable and nondurable goods. I first demonstrate that the low intertemporal elasticity of substitution of nondurables implies a very inelastic demand curve at a given point in time. I then show that the demand for durables has a much higher demand elasticity, reflecting the high substitutability of durables purchases over time.

Denote by γ the Lagrange multiplier on the accumulation equation of durable goods (the first equation in 1.2) and recall that λ is the Lagrange multiplier on the nominal budget constraint (1.1). Hence, γ is the shadow value of good X, expressed in utils, and λ represents the marginal utility of one dollar. Optimal behavior of the representative household (section 1.2.1.1) then implies that

$$\gamma_t = u_D\left(C_t, D_{H,t}\right) + \beta\left(1-\delta_D\right)\mathbb{E}_t\left[\gamma_{t+1}\right] \tag{1.14}$$

and

$$P_{X,t} = \lambda_t^{-1} \gamma_t. \tag{1.15}$$

As equation (1.14) shows, the shadow value of good X equals a flow component u_D plus a continuation value. Solving this equation forward and combining the result with (1.15) yields

$$P_{X,t} = \lambda_t^{-1} \mathbb{E}_t \sum_{s=0}^{\infty} \left[\beta \left(1 - \delta_D\right)\right]^s u_D \left(C_{t+s}, D_{H,t+s}\right). \tag{1.16}$$

This expression can be interpreted as the household's (inverse) demand function for X. It is helpful to first discuss the special case in which $\delta_D = 1$.

1.2.2.1 Nondurable goods

If $\delta_D = 1$, the small sector produces a nondurable good, $D_{H,t} = X_{H,t}$, and equation (1.16) reduces to the familiar expression

$$P_{X,t} = \lambda_t^{-1} u_D \left(C_t, X_{H,t}\right). \tag{1.17}$$

Next, define the intertemporal elasticity of substitution of D_H as $\sigma_D = -\frac{u_D(C,D_H)}{D_H u_{DD}(C,D_H)}$. The linear approximation of equation (1.17) then implies that, ceteris paribus, a one percent change in $X_{H,t}$ reduces $P_{X,t}$ by σ_D^{-1} percent. Hence, a low intertemporal elasticity of substitution σ_D implies that the demand curve for nondurable goods is inelastic.

Estimates of elasticity σ_D vary and are somewhat controversial. A recent study by *Cashin and Unayama* (2012) which explicitly distinguishes nondurable from storable and durable goods estimates a value 0.21, similar to *Hall*'s (1988) estimates. Such a low value suggests that consumption barely responds to changes in intertemporal prices or, equivalently, that the demand curve for nondurables is very steep. Even if σ_D is set to unity, a popular choice in the literature, the demand for nondurables is much less elastic than that for durables.

1.2.2.2 Durable goods

Due to the dynamic nature of the model, an analogous elasticity for durable goods ($\delta_D < 1$) is difficult to obtain. Instead, I will argue on the basis of equation (1.16) that in a limiting case the shadow value of the durable good γ_t is approximately constant. If γ_t is constant, the demand for X_t is perfectly elastic (see equation 1.15). The limiting approximation assumes that β approaches unity, δ_D approaches zero and that all disturbances are short-lived — a reasonable assumption for temporary fiscal expansions. Of course, many durable goods do not have depreciation rates near zero. For instance, the Bureau of Economic Analysis (BEA) estimates the service life of furniture at 14 years.[6] Yet, the approximation provides the correct intuition that the demand elasticity rises with greater durability and it remains quite accurate for realistic calibrations of δ_D. I provide details on the numerical accuracy of this approximation below.

To see why γ_t is approximately constant for long-lived durables, first notice that the consumer derives utility from the stock of the durable good $D_{H,t}$, not current purchases $X_{H,t}$. Because durables with long service lives have large stock to flow ratios (in steady state $D_H/X_H = 1/\delta_D$), even large changes in $X_{H,t}$ cause only relatively small percentage changes in $D_{H,t}$.

Second, if the household is sufficiently patient (β close to unity) and the durable long-lived (δ_D close to zero), the shadow value γ_t depends on utility flows far in the future. In stationary environments with short-lived shocks these future terms are barely affected since the economy quickly reverts back to its steady state. Any changes to the first few terms in the sum of equation (1.16) are dwarfed by the future terms which remain approximately unchanged. Hence, as β approaches unity and δ_D approaches zero, the shadow value γ_t becomes unresponsive to temporary shocks and the demand for $X_{H,t}$ perfectly elastic.[7]

[6]For the BEA's estimates of service lives, see *Bureau of Economic Analysis* (undated).
[7]A similar argument is made in *Barsky et al.* (2007).

Intuitively, the household smoothes consumption of both nondurable and durable goods. But since utility is derived from the stock the consumer is willing to tolerate much larger variation in durables purchases than in nondurables purchases. It is optimal to purchase durables only at favorable prices and to draw down the stock whenever prices are temporarily high. *Mankiw* (1985), *Adda and Cooper* (2000), *Erceg and Levin* (2006), *Mian and Sufi* (2012) and *Hausman* (2015) all provide evidence for large intertemporal substitutability of durable goods purchases.

1.2.3 Implications for fiscal policy

The demand elasticity is crucial for the size of the fiscal multiplier because it determines the degree to which private sector spending is crowded out. The basic intuition can then be illustrated in a demand and supply diagram.

Suppose the government raises spending and shifts out the demand curve. If the supply curve is upward-sloping, the resulting price increase reduces private sector spending. Greater demand elasticities lead to greater crowding-out. In the limiting case with horizontal demand curve, all private sector spending is crowded out and the fiscal expansion has no effect. Figure 1.1 illustrates the effect of government spending in a sector with inelastic demand (Panel A) and elastic demand (Panel B). The supply curve and the fiscal expansion are the same in both sectors.

Before turning to a more formal analysis of fiscal policy I briefly define sector-level analogues of the aggregate fiscal multiplier $\mathrm{d}Y/\,\mathrm{d}G$. These sectoral multipliers help link the theoretical analysis that follows to the empirical evidence in Section 1.3.

In the model described above, gross output in constant prices in sector X is $GO_{X,t} = P_X X_t$ and value added is $VA_{X,t} = P_X X_t - P_Z M_{X,t}$. Also denote government purchases of good X in constant prices by $G_{X,t} = P_X X_{G,t}$. Then the gross output and value added multipliers for sector X are defined as $\mathrm{d}GO_{X,t}/\,\mathrm{d}G_{X,t}$ and $\mathrm{d}VA_{X,t}/\,\mathrm{d}G_{X,t}$. Since I will estimate these multipliers in Section 1.3, it is of interest how these sector-level multipliers map into the policy relevant aggregate multiplier. When discussing the results below I pay particular attention to this relationship.

I next turn to two limiting approximations which allow me to solve for durable and nondurable multipliers analytically. The approximations require assumptions similar to those made above: I assume that β approaches unity and that the depreciation rate of capital δ_K tends to zero. The economy is then shocked with a short-lived increase in government purchases, X_G, in the small sector. To preserve space I limit myself to the discussion of these results and provide details on the derivations in Appendix A.1.2.

1.2.3.1 Spending on durable goods

I first consider the case in which the small sector produces highly durable goods.

Approximation result 1. *Suppose δ_K and δ_D are arbitrarily close to zero and β is arbitrarily close to* 1*. Then, for a short-lived increase in spending, it is approximately true that (1) $\Delta X_{H,t} \approx -\Delta X_{G,t}$, (2) the price $P_{X,t}$ remains unchanged, (3) the sectoral multipliers for gross output and value added are zero, $\frac{\mathrm{d}GO_{X,t}}{\mathrm{d}G_{X,t}} \approx \frac{\mathrm{d}VA_{X,t}}{\mathrm{d}G_{X,t}} \approx 0$, and (4) the aggregate multiplier is zero, $\frac{\mathrm{d}Y_t}{\mathrm{d}G_{X,t}} \approx 0$.*

Here Δ denotes the absolute deviation of a variable from steady state. Notice first that this result only requires the stated parametric assumptions. In particular, the result holds regardless of the functional forms of u and v and regardless of whether the X sector requires intermediates for production. The result is also independent of the degree of price stickiness (as long as prices are not perfectly sticky), the relative sizes of the two sectors, and the precise specification of the monetary policy rule (equation 1.13).

Part (1) of the result states that every dollar spent by the government crowds out one dollar of private sector spending. If prices are not perfectly sticky, the small sector's supply curve is upward-sloping. Hence, a greater quantity would lead to a higher equilibrium price. In this limiting approximation, however, the private sector's demand for durable goods is perfectly elastic and any price increase would result in complete withdrawal from the market. Clearly, this cannot be an equilibrium outcome. Instead, the equilibrium quantity and price in the small sector both remain unchanged. This requires that private sector demand contracts dollar for dollar with greater government spending.

Since the equilibrium quantity in the small sector is unaffected by the fiscal expansion, the sectoral multipliers must be zero. It turns out that in this limiting case the aggregate multiplier is zero as well. The explanation of this result has three components. First, as noted above, the fiscal expansion has no effect on the small sector's output and therefore leaves factor demands unchanged. It follows that the household's labor and capital income remain the same and that there are no spillovers to the large sector through the demand for intermediate inputs. The second effect concerns the government's financing of the rise in spending. Because the expansion is by assumption brief, its effect on life-time income through taxation is very small. In fact, the approximation procedure treats the change in life-time income as negligible. It then follows that households neither change their labor supply nor their overall consumption demand. Finally, notice that both prices P_X and P_Z remain unchanged. Hence greater government spending does not raise inflation or output and no adjustment is required for the nominal interest rate.

An important corollary of this approximation result is that fiscal multipliers need not be large at the zero lower bound (ZLB). As, among others, *Christiano et al.* (2011) and *Woodford* (2011) show, the multiplier is large whenever the fiscal expansion leads to inflation, and these inflationary forces are not offset by a higher policy rate. At the ZLB, higher inflation reduces the real rate, stimulating private consumption and therefore resulting in a large fiscal multiplier. However, when the government purchases highly durable goods, private spending is crowded out and inflation barely rises. The mechanism of greater demand leading to greater inflation and greater inflation leading to even greater demand is therefore not triggered. As a result, the durables multiplier remains low.

While the multipliers for long-lived durable goods are small at the ZLB, this is not always true for durable goods with intermediate service lives. I return to this issue in Section 1.4 in which I provide a quantitative analysis of durables multipliers at the ZLB.

1.2.3.2 Spending on nondurable goods

I next turn to the case in which the small sector produces a nondurable good.

Approximation result 2. *Suppose $\delta_D = 1$, δ_K is arbitrarily close to zero and β is arbitrarily close to 1. Suppose further that prices are fully flexible, and that $u_{CD} = v_{XZ} = 0$. Lastly, there is an additional technical assumption of little economic relevance which I discuss in Appendix A.1.2. Then a short-lived increase in spending yields a gross output sectoral multiplier equal to*

$$\frac{\mathrm{d}GO_{X,t}}{\mathrm{d}G_{X,t}} \approx \frac{\left(1+\eta_X^{-1}\right)\chi^{-1} - \left(\alpha+\eta_X^{-1}\right)}{\sigma_D\left(\alpha+\eta_X^{-1}\right)\frac{X_H}{X} + \left(1+\eta_X^{-1}\right)\chi^{-1} - \left(\alpha+\eta_X^{-1}\right)}$$

and approximately equal sectoral value added and aggregate multipliers

$$\frac{\mathrm{d}VA_{X,t}}{\mathrm{d}G_{X,t}} \approx \frac{\mathrm{d}Y_t}{\mathrm{d}G_{X,t}} \approx \frac{1-\alpha}{\sigma_D\left(\alpha+\eta_X^{-1}\right)\frac{X_H}{X} + \left(1+\eta_X^{-1}\right)\chi^{-1} - \left(\alpha+\eta_X^{-1}\right)}.$$

Here, $\eta_X = \frac{v_X(N_Z,N_X)}{N_X v_{XX}(N_Z,N_X)}$. The relative price $P_{X,t}/P_{Z,t}$ rises in response to greater spending.

When the small sector produces a nondurable good, all three multipliers are positive. More precisely, they are bounded between zero and one, a feature common in neoclassical environments. Additionally, the sectoral value added multiplier and the aggregate multiplier are again approximately equal, implying that sectoral multipliers are of direct policy interest.

I next discuss how various parameters affect the multipliers. The main property that this paper emphasizes is the multipliers' dependence on the intertemporal elasticity of substitution σ_D. The greater σ_D the smaller the multipliers. As σ_D approaches infinity all three multipliers tend to zero, the same value that multipliers of highly durable goods take.

To understand the role of the remaining parameters, I consider two polar cases. Suppose first that no intermediates are required for production ($\chi = 1$). Then it is easy to show that all three multipliers approximately equal

$$\frac{1-\alpha}{1-\alpha+\sigma_D\left(\alpha+\eta_X^{-1}\right)\frac{X_H}{X}}.$$

This formula is identical to that in *Hall* (2009, p. 199). It illustrates clearly that the multiplier is increasing in the labor supply elasticity, η_X, and decreasing in the capital share, α, reflecting the fact that capital is a fixed factor in the short run.

I next turn to the opposite case in which $\chi \to 0$ so that the small sector almost exclusively uses intermediates in production. It then follows that the sectoral multiplier for gross output, $\frac{\mathrm{d}GO_{X,t}}{\mathrm{d}G_{X,t}}$, approaches unity while the sectoral multiplier for value added and the aggregate multiplier tend to zero. What is the intuition behind these results?

A gross output multiplier of unity implies that the small sector expands one-for-one with greater government demand. Since additional output is almost exclusively produced from intermediates, it is clear that value added in the small sector remains close to zero. It can also be shown that

$$\Delta M_{X,t} \approx -\left(\Delta I_{X,t} + \Delta I_{Z,t}\right),$$

so that purchases of intermediates crowd out investment in the large hybrid sector dollar for dollar. Hence, while the small sector expands one-for-one, this expansion has no effect on production in the large sector and total value added (GDP) remains unchanged. We therefore encounter a second instance in which the crowding out of durable goods with large intertemporal elasticity of substitution—in this case investment goods—implies a low multiplier.

Although this finding is certainly extreme and crucially relies on the assumption that the large sector's output can be used for investment, the approximation clearly illustrates how sectoral linkages affect fiscal multipliers: If government purchases, either directly or indirectly through intermediate input linkages, raise the demand for durable goods, they largely crowd out private sector demand. If, in contrast, government purchases are targeted at nondurable goods, there is less crowding out and the multipliers are larger.

1.2.4 Numerical results

I next use an exact (linear) solution of the model to confirm that the durables multiplier is relatively small. Before proceeding, however, I modify the model slightly. In order to avoid the extreme crowding out effects that greater intermediate purchases cause in the large hybrid sector, I introduce investment adjustment costs. These adjustment costs reflect the fact that some intermediates are nondurable in nature and therefore have a less elastic demand function than those intermediates that are durable.

1.2.4.1 Calibration

The length of a period is a quarter and households discount the future with discount factor $\beta = 0.99$. The flow utility function is given by

$$u\left(C_t, D_{H,t}\right) = \left(1 - \frac{1}{\sigma}\right)^{-1} \left(\left[\omega \left(C_t\right)^{\frac{\rho-1}{\rho}} + \left(1-\omega\right)\left(D_{H,t}\right)^{\frac{\rho-1}{\rho}}\right]^{\frac{\rho}{\rho-1}}\right)^{1-\frac{1}{\sigma}}. \tag{1.18}$$

Consistent with the estimates in *Hall* (1988) and *Cashin and Unayama* (2012), I select an intertemporal elasticity of substitution σ of 0.25. In the baseline calibration, I further assume that the elasticity of substitution ρ between C and D_H is (arbitrarily close to) 1. I explore alternative values in robustness exercises. The preference weight ω is chosen so that value added in the small sector is one hundredth of total value added.

I assume that the disutility of labor is given by

$$v\left(N_{X,t}, N_{Z,t}\right) = \varphi_X \frac{\left(N_{X,t}\right)^{1+\frac{1}{\eta}}}{1+\frac{1}{\eta}} + \varphi_Z \frac{\left(N_{Z,t}\right)^{1+\frac{1}{\eta}}}{1+\frac{1}{\eta}}.$$

This specification implies that labor is immobile across sectors, an assumption I also relax in robustness exercises. I set the labor supply elasticity η to unity. This value is broadly consistent with recent suggestions in the literature *Kimball and Shapiro* 2008, *Hall* 2009, and *Chetty et al.* 2011).[8]

Turning to the production side of the model, I choose $\alpha = 1/3$ as is standard in the literature. Consistent with an intermediate input share of roughly 55 percent in the 2007 Make and Use Tables, I calibrate χ to 0.45. Further, I set ε to 6, implying a steady state mark-up of 20 percent. In my choice of the the price stickiness parameters θ_X and θ_Z, I follow *Galí* (2008) and assume that both equal 2/3. The depreciation rates of both types of capital are set to 0.025, implying approximately a 10 percent annual depreciation rate. I show results for various alternative values of δ_D, including $\delta_D = 1$ so that the small sector produces a nondurable good.

In steady state, the government purchases one fifth of the output in the small sector. I calibrate the persistence ϱ_X of the spending process (1.10) to 0.75. This value implies that the fiscal expansion largely dissipates after eight quarters, mimicking the spending trajectory of the American Recovery and Reinvestment Act.[9] The monetary authority follows a simple Taylor rule, that is, I replace the general rule (1.13) by $i_t = \beta^{-1} - 1 + \phi_\pi \pi_t + \phi_Y \tilde{Y}_t$. Here,

[8] Once ω and the remaining parameters are fixed, the values of φ_X and φ_Z do not affect equilibrium dynamics.

[9] The Congressional Budget Office estimates that 78 percent of total spending had occurred by September 2011 (see `https://www.cbo.gov/publication/42682`).

the tilde denotes percent deviation from steady state. In the baseline calibration I set $\phi_\pi = 1.1$ and $\phi_Y = 0$. Notice that this policy rule satisfies the Taylor principle, ensuring equilibrium determinacy as shown in *Bullard and Mitra* (2002). I consider a more hawkish parameterization of the Taylor rule in Appendix A.1.3.

Finally, I assume that capital investment at time t is subject to adjustment costs of the form

$$K_{j,t-1}\frac{\zeta_K}{2}\left(\frac{I_{j,t}}{K_{j,t-1}} - \delta_K\right)^2, \quad j \in \{X, Z\}.$$

In the baseline calibration, summarized in Table 1.1, ζ_K is set to 20. This value is slightly higher than the 17 chosen by *Christiano et al.* (2011) who base their calibration on *Eberly et al.* (2008).

1.2.4.2 Simulation results

I next turn to the question how the economy's reaction to a government spending shock depends on the depreciation rate δ_D. I consider four different cases. First, I set $\delta_D = 1$ so that the small sector produces nondurables. Second, I set $\delta_D = 0.083$, a value consistent with a 3 year service life. The *Bureau of Economic Analysis* (undated) estimates for example the service lives of tires and software at 3 years. Third, I select a value of $\delta_D = 0.018$. This value corresponds to a service life of 14 years (e.g. furniture). Lastly, I consider a very long-lived durable. According to the BEA, new 1-to-4-unit structures have service lives of 80 years, implying a quarterly depreciation rate of approximately $\delta_D = 0.003$. Figure 1.2 shows the impulse response functions of a 100 unit increase of government spending in the X sector.

Panel A displays the fiscal expansion in the public sector and Panel B shows the private sector response. If the small sector produces a nondurable good, private sector spending barely changes. Since there is very little crowding out, production expands almost one-for-one with public spending. This is shown in Panel C. The aggregate multiplier is near 0.75

(Panel D) and the gross multiplier in the X sector is close to unity (Panel E). If, in contrast, the small sector produces durable goods, the multipliers are much smaller.

The lower the depreciation rate the more private sector spending is crowded out. When the X sector produces a moderately durable good with a 3 year service life ($\delta_D = 0.083$), private sector spending falls substantially more after the spending shock. All multipliers, gross output, value added, and aggregate, fall to roughly two thirds of the nondurable goods multiplier. For a service life of 14 years, the multipliers fall to about 1/3 of the nondurables multiplier. As the depreciation rate further approaches zero the multipliers gradually shrink to zero. The model therefore predicts that infrastructure investment which crowds out private sector construction has a very small fiscal multiplier of roughly 0.1.

These impulse response functions corroborate the results presented earlier: The demand for durable goods is highly elastic and easily crowded out. However, they yield one additional insight. Because all real variables of the model are stationary, they return to their steady state values in the long run. This model feature implies that any demand that is temporarily crowded out has to be made up in the long run as households replenish their stock of durables. In fact, Panel B shows that households' purchases of the durable goods rise above zero roughly 10 quarters after the shock.

An implication of this delayed demand boost is that the multipliers of durable goods begin to rise about 4 quarters after the shock. From the viewpoint of stabilization policy, it is an unfortunate fact that private demand picks up only after the fiscal expansion ends and the pressure on prices recedes. For short-lived spending shocks the economy will never enjoy increased government and private sector demand at the same time. I report the impulse response functions of the price in sector X and the nominal interest rate in Figure A.1 in Appendix A.1.4.

I finally turn to the question of what sectoral multipliers teach us about aggregate multipliers. In the limiting approximations above, crowding out of capital investment was perfect and, as a result, the aggregate multipliers were equal to the sectoral value-added multipliers. However, when the depreciation rate of capital is higher and investment is subject to adjustment costs, as in the calibrations shown here, crowding out is imperfect. The aggregate multiplier is now greater than the sectoral value added multiplier. In fact, for all four calibrations shown in Figure 1.2, the aggregate multipliers lie above the sectoral value added multipliers and are quite similar to or somewhat below the gross output multipliers. As I show in Appendix A.1.3, the rule of thumb that the aggregate multipliers are slightly smaller than the sectoral gross output multipliers is robust for a number of alternative calibrations.

1.3 Empirical evidence

The main objective in this section is to test whether the data support the hypothesis that the durables multiplier is smaller than the nondurables multiplier. I begin the analysis using national accounts data and find that the evidence supports the hypothesis. I then turn to an analysis at the industry level and again find evidence for smaller multipliers of durable goods purchases.

1.3.1 Evidence based on aggregate data

For the analysis of aggregate data, I adopt a specification similar to *Hall* (2009) and *Barro and Redlick* (2011), namely

$$\frac{Y_t - Y_{t-1}}{Y_{t-1}^T} = \alpha + \mu_X \frac{G_{X,t} - G_{X,t-1}}{Y_{t-1}^T} + \mu_C \frac{G_{C,t} - G_{C,t-1}}{Y_{t-1}^T} + \nu RN_t + \varepsilon_{t+1}. \tag{1.19}$$

In this equation Y_t is GDP, Y_t^T is trend GDP, $G_{X,t}$ is military spending on durable goods, $G_{C,t}$ is military spending on nondurable goods, and RN_t is *Ramey*'s (2011) defense news variable. Ramey's news variable is constructed from narrative records and measures the present value of new military spending at the time of announcement. The variable mostly captures defense spending 3 to 5 years into the future and is expressed as a fraction of GDP.

I control for the announcement of future spending to isolate the effect of actual spending on durable and nondurable goods. According to standard theory, the announcement of greater future spending reduces households' life-time wealth and increases their labor supply. ν is therefore expected to be positive. The main reason for including this control is that announcements of future spending could be correlated with one of the spending variables but not the other. If, for example, the public learns about a military buildup at the same time as first investments into equipment are made, the estimate of μ_X would be biased upward without the control for news on future spending. Note that when μ_X and μ_C are estimated based on specification (1.19), these "multipliers" are purged of announcement effects and thus understate the total effect of military spending.[10]

Consistent estimation of the multipliers μ_X and μ_C requires the commonly made assumption that the state of the business cycle does not affect whether the U.S. engages in military buildup (e.g. *Ramey and Shapiro* 1998, *Hall* 2009). I use annual national accounts data for the estimation. Details are available in Appendix A.2.1.

Table 1.2 shows the results. The output multipliers are quite different for durable and nondurable goods. For the sample from 1929 to 2014 the estimate of μ_C is 0.49. When World War II is excluded, the multiplier rises to 0.83. In contrast, the estimates of μ_X are close to zero or negative, supporting the theoretical predictions that durables multipliers are smaller than nondurables multipliers. Yet, the standard errors are too large to reject the null hypothesis that μ_C equals μ_X. As expected, the estimates of ν are positive.

I next replace GDP in the numerator on the left hand side of equation (1.19) with a number of other variables to estimate how government spending on durable and nondurable goods crowds out private sector spending. These are consumption, investment, consumption

[10] Based on the argument in *Barro and Redlick* (2011) the total size of the durables and nondurables multipliers should roughly be $\mu_X + 4 \cdot \nu$ and $\mu_C + 4 \cdot \nu$.

of nondurables and services, investment plus durable goods consumption, and net exports. It is striking how consistently the coefficient on nondurables is estimated to be essentially zero. Hence, there is little, if any, crowding out when the government purchases nondurable goods. There is, however, crowding out when the government purchases durable goods. For the sample from 1929 to 2014 the estimate of μ_X is significantly negative for all four measures of private consumption and investment. The coefficients are near negative one for investment and investment plus durables consumption, suggesting that there is almost perfect crowding out. It is less clear how the negative coefficient on durables spending in the consumption equation should be interpreted. When consumption is limited to nondurables and services, the estimates of μ_X move slightly towards zero. Although this suggests that there is indeed more crowding out of durables consumption, the estimates of μ_X remain negative.

The model in section 1.2 suggests that the durables multiplier is small because the private sector substitutes intertemporally when the government raises spending. An alternative explanation is that the private sector substitutes towards imported goods. If this was the case, greater government spending should be accompanied by lower net exports and a smaller drop in investment. The last line in Table 1.2 shows that the response of net exports is negligibly small. Hence, the data do not support the alternative hypothesis that the durables multiplier is small because of substitution towards foreign goods.

From 1972 onwards, the BEA provides a more detailed breakdown of military spending. Evidence based on this data again suggests that fiscal policy is less effective when the government purchases durable goods. I discuss data, estimation strategy and results in Appendix A.2.2 and next turn to evidence based on industry-level data.

1.3.2 Evidence based on industry-level data

In this section I estimate fiscal multipliers at the industry level: How much does industry output rise if the government spends 1$ on goods in this industry? Moving towards less aggregate data has the benefit of substantially raising the available data and therefore statistical power. As I will discuss below, this analysis has the additional advantage of relaxing the identifying assumptions. Yet, there is a mild cost: The policy relevant aggregate multipliers generally differ somewhat from the estimated industry-level multipliers. Since I established a close connection between these two types of multipliers in section 1.2, I view these costs as relatively small.

1.3.2.1 Data

The empirical analysis at the industry level is based primarily on two data sources, the NBER-CES Manufacturing Industry Database and the military prime contract files. The NBER-CES database contains annual data on, among other things, gross output, value added, cost of materials, expenditures on energy, and employment, along with various deflators. It ranges from 1958 to 2009 and covers all manufacturing industries. The database is constructed mainly from the Annual Survey of Manufactures and the Census of Manufactures, but complemented with additional information from the Bureau of Economic Analysis, the Bureau of Labor Statistics, and the Federal Reserve Board. A detailed description of this database is provided by *Bartelsman and Gray* (1996) and *Becker et al.* (2013).

The military prime contract files include information on all military prime contracts with values above the minimum threshold of $10,000 up to 1983 and $25,000 thereafter. They can be downloaded for the period from 1966 to 2003 from the U.S. National Archives. I complement the prime contract files with data from USAspending.gov, a government website dedicated to promoting transparency of federal spending. The data from USAspending.gov is available from 2000 onwards. A comparison of the two data sources for the overlapping years from 2000 to 2003 reveals only negligible differences. The analysis below is based on all contracts that are awarded to firms in the United States.

Unfortunately, the data on defense spending is not easily matched to different industries.

While the NBER-CES database is available for both SIC- and NAICS-based industry definitions, the military prime contract files contain SIC codes only for the relatively brief period from 1989 to 2000 and NAICS codes from 2000 onwards. Instead, military purchases are classified according to the Federal Procurement Data System which assigns a unique Product Service Code (PSC) or Federal Supply Code (FSC) to each contract since 1966.

To obtain military spending at the industry level I construct a concordance from the FSC/PSC classification to 4-digit SIC codes. The concordance is based on the military prime contract files from 1989 to 2000 which contain both, FSC/PSC and SIC codes. Details on the construction of this concordance as well as further information on the data and the FSC/PSC classification system are available in Appendix A.3.1. Because the FSC/PSC system underwent a major revision in 1979, the concordance is only valid thereafter. This leaves me with a sample of annual data from 1979 to 2009. I adopt the SIC classification of durable and nondurable goods.

Due to concerns about measurement error (see below) and because one would not expect that small changes in military spending give rise to measurable changes in economic activity, I limit the sample to industries in which the military purchases at least 1 percent of gross output, on average. Examples of dropped industries include Greeting Cards (SIC 2771) and Women's Footwear (SIC 3144). I also drop industries with little private sector demand. These industries are problematic for testing the hypothesis of smaller multipliers for spending on durable goods because with little private sector demand to begin with, there is little room for crowding out. In the complete absence of private sector demand, the theory discussed above does not apply and the sectoral multiplier should be unity regardless of whether the sector produces durable or nondurable goods. I therefore drop industries with average values of military purchases per industry gross output of greater than 0.35.[11] Examples here include Tanks and Tank Components (SIC 3795) and Ammunition, except for Small Arms

[11] Other thresholds leave the results essentially unchanged.

(SIC 3483). The final sample comprises 35 nondurables and 76 durables industries. I list all industries in the sample in Appendix A.3.2.

1.3.2.2 Empirical strategy

Specification

I first estimate impulse response functions using *Òscar Jordà*'s (2005) local projection method and then construct multipliers from the estimated impulse response functions. The baseline specification is

$$\begin{aligned}\frac{Y_{i,t+h} - Y^T_{i,t-1}}{VA^T_{i,t-1}} \quad = \quad & \alpha_h \frac{G_{i,t} - G^T_{i,t-1}}{VA^T_{i,t-1}} + \sum_{k=1}^{2} \beta^k_h \frac{Y_{i,t-k} - Y^T_{i,t-k-1}}{VA^T_{i,t-k-1}} \\ & + \sum_{k=1}^{2} \gamma^k_h \frac{G_{i,t-k} - G^T_{i,t-k-1}}{VA^T_{i,t-k-1}} + \delta_{i,h} + \zeta_{t,h} + \varepsilon_{i,t+h}, \end{aligned} \tag{1.20}$$

for $h = 0, 1, ..., 4$. In this equation $Y_{i,t}$ is a generic variable of interest of industry i at time t, $G_{i,t}$ is defense spending, and $VA_{i,t}$ is value added. The superscript T indicates that the variable in question is an HP-filtered trend.

I estimate equation (1.20) separately for durable and nondurable goods industries to obtain the objects of interest $\{\alpha_h\}^4_{h=0}$. These parameters represent the impulse response coefficients for the impact year, $h = 0$, and four subsequent years. The specification controls for two lags of the deviations of the dependent variable and defense spending from trend, as well as time and industry fixed effects $\delta_{i,h}$ and $\zeta_{t,h}$.

The time fixed effects play a key role in specification (1.20). They soak up disturbances that affect all industries symmetrically, notably monetary policy shocks and certain tax policy changes. Additionally, the time fixed effects control for announcements of greater future spending, at least to the extend that all sectors are affected equally. Notice that since I estimate equation (1.20) separately for durable and nondurable goods industries all coefficients are allowed to differ by industry type.

A potential concern with specification (1.20) is that the impulse response coefficients $\{\alpha_h\}_{h=0}^4$ depend on the smoothness of the trends required for constructing the left and right-hand side variables. To err on the safe side I extract very smooth trends with a smoothing parameter of 1600 for annual data. I also test the robustness of my results for alternative values of this parameter.

Identification

As *Nekarda and Ramey* (2011) discuss in detail, an industry-level analysis of government spending may suffer from an endogeneity problem. Technological progress in a particular industry can lead both to greater sales to the private sector and to increased defense spending as the military upgrades its equipment. The bias resulting from the estimation of (1.20) by OLS would inflate the impulse response coefficients above their true levels.

On the other hand, measurement error in defense spending may bias the coefficients towards zero. Measurement error is a concern for the following reason. In the military prime contract files, every contract is assigned a single FSC/PSC code. The documentation of the files reveals occasional difficulties of the procurement staff to select a single code, as some contracts include purchases of different types of goods.[12] Although the staff is instructed to assign the FSC/PSC code whose description best fits the contract, it is likely that the limitation to one code per contract induces measurement error. A second source of error is the use of the concordance to map FSC/PSC codes to SIC industry definitions.

To avoid or reduce these biases I construct a Bartik-type instrument from total defense spending, i.e. spending summed over all manufacturing industries (*Bartik* 1991). More

[12] For example, guns, up to 30mm (FSC 1005) and guns, over 30mm up to 75mm (FSC 1010) are two separate categories. It is conceivable that purchases of both types of guns were part of the same contract.

precisely, I construct the variable

$$\frac{1}{5}\sum_{s=1979}^{1983}\frac{G_{i,s}/GO_{i,s}}{G_s/GO_s}\cdot\frac{G_t-G_{t-1}^T}{VA_{t-1}^T}. \tag{1.21}$$

The first term in this expression is a five-year average of the industry-specific military spending share divided by the aggregate military spending share. It scales the aggregate military spending series (the second term) so as to generate industry-specific variation. The idea is that those industries in which the military purchases a greater fraction of output, on average, also experience greater changes in spending in response to aggregate military buildups or drawdowns.[13] Because all military spending variables on the right-hand side of baseline specification (1.20) are potentially endogenous, I use (1.21) and its two lags as instruments.

Two assumptions are required for (1.21) to satisfy the exclusion restriction. First, and recalling that specification (1.20) contains time fixed effects, the *relative* performance of industries does not affect whether the U.S. government engages in an aggregate military buildup. This assumption rules out a reverse causality problem of the type discussed above. Second, it is not the case that aggregate military spending changes other determinants of industry-level output after controlling for the right-hand side variables of specification (1.20). As noted above, monetary and tax policy as well as a nationwide draft do not pose a threat to identification because the effects should be roughly symmetric within durable or nondurables industries and will therefore be soaked up by the time fixed effect. A concern would arise, however, if the cyclicality in the absence of military spending was greater in industries which receive greater fractions of government spending. While this condition is inherently not testable, it is reassuring to note that the average standard deviation of value added growth in industries which receive below and above median military spending are very similar.

I next turn to the first stage of the estimation. Table 1.3 summarizes the Angrist-Pischke

[13] *Nekarda and Ramey* (2011) and *Nakamura and Steinsson* (2014) use similar approaches to construct instruments.

F-statistics of excluded instruments and their p-values when the dependent variable is value added for the time horizons from $h = 0, ..., 4$. The smallest F-statistic in Table 1.3 takes the value 15, implying that the instruments are strong. In principle, the first stage for specification (1.20) is different for every dependent variable $Y_{i,t}$ (due to the lags on the right-hand side). In practice, however, the Angrist-Pischke F-statistics are virtually identical. I therefore only report the details for value added.

1.3.2.3 Results

Figure 1.3 shows the estimated impulse response functions for the baseline sample. A unit increase in military spending leads to additional spending in subsequent years (Panel A). The impulse response functions for both types of industries begin to fall in year five and return to zero 6 to 7 years after the shock (see Figure A.2, Appendix A.3.3). Hence, the spending increase is more persistent than a typical stimulus program such as the ARRA. I show in Appendix A.1.3 that the model predicts a difference of factor two between the nondurables and the durables multiplier for shocks with this persistence.

Panel B shows the dynamic responses of gross output associated with these spending paths. Consistent with the hypothesis of greater crowding out in durable goods industries, the effect of military spending on gross output is quite small. In contrast, nondurables industries expand substantially.

Panel C displays the responses of value added. In both types of industries value added rises significantly above zero and, again, the dynamic response for durables industries lies below that for nondurables industries. Additionally, in nondurables industries the rise in valued added is accompanied by increased purchases of materials (Panel D), although the standard errors are fairly large. By contrast, the costs of materials change little in durables industries.[14] Panel E shows the impulse response functions of energy expenditures. Unfor-

[14]Value added and cost of materials roughly sum to gross output.

tunately, the standard errors are too large to allow for a conclusive statistical comparison. Finally, Panel F shows the employment responses. Consistent with theory, employment rises substantially more in nondurable goods industries.[15]

For the interpretation of the impulse responses in the impact period note that military spending by industry is constructed by aggregating the value of all contracts in a given year. I use the date on which the contract is signed for this aggregation. My dataset has no information on the date of actual payments. Panels B to F all suggest that there is little effect at the time the contract is signed, but only in subsequent years.

Taken together, nondurable goods industries respond strongly to increased defense spending while the reaction of durable goods industries is quite moderate. These findings are consistent with the theory described above, suggesting that indeed there is little crowding out in nondurable goods industries but substantial crowding out in durables industries.

I next compute the sectoral multipliers as the cumulative change in the outcome variable divided by the cumulative change in spending. Based on the model from the previous section and taking into account that military buildups typically have greater persistence than stabilizing interventions, I expect that the durables multiplier has about half the size of the nondurables multiplier.

Table 1.4 summarizes the estimated multipliers for time horizons of 1 to 3 years after the shock. The table confirms that multipliers for all five variables, gross output, value added, cost of materials, energy expenditures, and employment are uniformly larger in nondurable goods sectors. Notice that the employment multiplier is expressed as employees per year per $1 million dollar of spending. While most multipliers in Table 1.4 take empirically plausible sizes and are broadly in line with theoretical predictions, the gross output and cost of materials multipliers for nondurables are quite large. A likely explanation for this is that measurement problems give rise to upward biases. If firms in a particular industry use

[15] I report estimates of impulse response functions for prices in Appendix A.3.5.

inputs from other firms in the same industry, shipments and cost of materials will be counted multiple times. The Use Tables of various years suggest that 10 percent is a conservative estimate of the share of intra-industry shipments. Under this assumption, one dollar of final sales is counted $1/(1-0.1) \approx 1.11$ times. A similar problem arises from intra-firm, inter-plant shipments.[16] For this reason it is preferable to compare industries on the basis of value added or employment.[17]

Robustness

I next discuss the robustness of these results. All estimates are reported in Appendix A.3.4.

One possible concern is that is that anticipation of future spending leads to asymmetric effects across industries which are not fully captured by time the fixed effects. To address this issue I add *Ramey*'s 2011 defense news variable interacted with industry indicators to specification (1.20). The resulting multipliers are almost identical to those in the baseline specification (Table A.7). A second concern may be that the time fixed effects do not fully control for monetary policy. The reason is that the interest sensitivity increases with the length of the service lives, and these vary across industries. When controlling for the real interest rate interacted with industry indicators, however, the results barely change (Table A.8).

I next explore how the results depend on how the trend of the variables is extracted. For the baseline results I used an HP-filter with a smoothing parameter of 1600. The estimates for the alternative smoothing parameters of 400 and 6000 are shown in Tables A.9 and A.10. For smaller values of the smoothing parameters the multipliers decrease slightly. Yet, there is not a single case in which a nondurables multiplier falls below the value of the corresponding

[16] *Atalay et al.* (2014) report that 16 percent of shipments occur within the firm.
[17] See *Chodorow-Reich et al.* (2012), for instance, for estimates of employment multipliers.

durables multiplier.

Finally, I estimate durables multipliers at the 3-digit SIC level.[18] This wider industry definition captures that spending on goods in one 4-digit industry may drive up prices of factors that are used in other 4-digit industries under the same 3-digit umbrella. As a result, one would expect larger crowding-out effects. Indeed, the durables multipliers at the 3-digit level are very close to zero (Table A.11).

1.3.2.4 The state of the business cycle

Several studies have argued that the fiscal multiplier is larger in slumps than in booms (e.g. *Auerbach and Gorodnichenko* 2012, 2013, *Michaillat* 2014). To see whether this form of state dependence applies to the durable goods multiplier, I estimate impulse response functions separately for slack and nonslack periods (which I somewhat imprecisely refer to as recessions and expansions). I adopt a specification similar to *Auerbach and Gorodnichenko* (2013),

$$\begin{aligned}\frac{Y_{i,t+h}-Y^T_{i,t-1}}{VA^T_{i,t-1}} = {} & \alpha^R_h F_{i,t-1}\frac{G_{i,t}-G^T_{i,t-1}}{VA^T_{i,t-1}} + \alpha^E_h\left(1-F_{i,t-1}\right)\frac{G_{i,t}-G^T_{i,t-1}}{VA^T_{i,t-1}} \\ & + \sum_{k=1}^{2}\beta^R_{h,k}F_{i,t-1}\frac{Y_{i,t-k}-Y^T_{i,t-k-1}}{VA^T_{i,t-k-1}} + \sum_{k=1}^{2}\beta^E_{h,k}\left(1-F_{i,t-1}\right)\frac{Y_{i,t-k}-Y^T_{i,t-k-1}}{VA^T_{i,t-k-1}} \\ & + \sum_{k=1}^{2}\gamma^R_{h,k}F_{i,t-1}\frac{G_{i,t-k}-G^T_{i,t-k-1}}{VA^T_{i,t-k-1}} + \sum_{k=1}^{2}\gamma^E_{h,k}\left(1-F_{i,t-1}\right)\frac{G_{i,t-k}-G^T_{i,t-k-1}}{VA^T_{i,t-k-1}} \\ & + \eta_h F_{i,t-1} + \delta_{i,h} + \zeta_{t,h} + \varepsilon_{i,t+h}. \end{aligned} \tag{1.22}$$

In this equation

$$F_{i,t} = \frac{\exp\left(-\kappa\cdot VA^C_{i,t}\right)}{1+\exp\left(-\kappa\cdot VA^C_{i,t}\right)},\ \kappa>0,$$

[18]The sample size for nondurables industries is too small to obtain informative estimates.

and $VA^C_{i,t}$ denotes the demeaned and standardized cycle component of value added in sector i. I use the one-sided HP-filter from *Stock and Watson* (1999) with a smoothing parameter of 1600 to extract the cycle component.[19][20] $F_{i,t}$ measures the "degree" to which industry i's value added is below trend (in recession). It varies between zero and one and takes greater values whenever the industry's value added is low. Hence, the empirical model (1.22) permits estimation of impulse response functions separately for recessions ($F_{i,t} = 1$) and expansions ($F_{i,t} = 0$). These impulse response functions are given by $\{\alpha^R_h\}^4_{h=0}$ and $\{\alpha^E_h\}^4_{h=0}$. Parameter κ is set to 1.5 which implies that the economy spends about 20 percent of the time in recessions, a value consistent with U.S. business cycle facts. For more details see *Auerbach and Gorodnichenko* (2012, 2013). Notice, that I include $F_{i,t-1}$ as a control variable. By doing so, I allow the left-hand side variable to directly depend on the state of the economy. To address endogeneity concerns, I construct instruments by multiplying (1.21) with $F_{i,t-1}$ and $1 - F_{i,t-1}$. The instruments are strong as the Angrist-Pischke F-statistics in Table show.

Figure 1.4 shows the impulse response functions starting with military spending in Panel A. A spending shock in recessions is followed by somewhat lower subsequent spending than a shock in expansions. There is little evidence for greater multipliers in recessions. The impulse response functions for gross output (Panel B), value added (Panel C), and cost of materials (Panel D) are initially negative in recessions and only gradually rise above zero. In expansions, these variables are positive at first and then return to values near zero. There is little information in the dynamic responses of energy expenditures (Panel E). Employment increases slightly after a fiscal shock in recessions, but after three years the impulse response function falls below the response in expansions (Panel F).

Table 1.5 shows the associated multipliers. They are often negative in recessions. Since the standard errors are large, this analysis cannot rule out that the durables fiscal multiplier

[19] I would like to thank Valerie Ramey and Simon Gilchrist for pointing out to me that it is critical to use a one-sided filter in this specification.

[20] *Meyer-Gohde* (2010) provides an implementation of the one-sided HP-filter by *Stock and Watson* (1999).

depends on the state of the economy. Yet, it is unlikely that the degree of state dependence is sufficiently strong to render the durable goods multiplier "large" in recessions.[21]

1.4 The zero lower bound

In this section I return to theory and analyze the sizes of durables and nondurables multipliers in an economy in which the zero lower bound (ZLB) on the nominal interest rate binds. As I showed in Section 1.2.3.1, for very low depreciation rates the multiplier for durable goods is small regardless of the monetary policy response. For moderate depreciation rates, however, the multiplier can be larger when the ZLB binds.

To study government spending at the ZLB, I modify the model from Section 1.2 in two ways. First, I replace the monetary policy rule (1.13) with

$$i_t = \max\{0, \beta^{-1} - 1 + \phi_\pi \pi_t + \phi_Y \tilde{Y}_t\},$$

where $\phi_\pi = 1.1$ and $\phi_Y = 0$. Second, I assume that the discount factor follows the AR(1) process

$$\beta_t = (1 - \varrho_\beta)\,\beta + \varrho_\beta \beta_{t-1} + \varepsilon_{\beta,t}. \tag{1.23}$$

The remaining model equations and the calibration remain unchanged.

I consider the following scenario. Prior to time 0, the economy is subject to a positive discount factor shock so that the ZLB begins to bind. The government then raises spending on X by 100 (artificial) quantity units at time 0. I choose the persistence of the discount factor shock ϱ_β to imply that the ZLB continues to bind for 4, 8, and 16 quarters beginning at time 0. The persistence of the spending shock is as in the baseline calibration, $\varrho_X = 0.75$. I assume that the government spending shock is sufficiently small to never lift the economy

[21] *Berger and Vavra* (2014) argue that the durable goods fiscal multiplier is smaller in recessions than in expansions.

out of the ZLB regime. The fiscal multiplier is then computed from the incremental output response to the shock in government spending.[22]

Figure 1.5 shows the multipliers together with the impulse response functions of the price level P. When the ZLB binds for 4 quarters (Panels A and B), the multiplier for nondurable goods is slightly below 2. The multiplier is smaller when the government purchases goods with greater durability. For durables with 10 year service lives the multiplier is below unity and for durables with 80 year service lives the multiplier is below 0.3. Hence, for a short period of 4 quarters at the ZLB the ordering of multipliers remains unchanged, but they are larger than when the central bank offsets the expansion by raising the interest rate.

At the ZLB greater government spending raises inflation and thereby lowers real interest rates. Lower real rates, in turn, crowd in private sector spending. Hence, to understand the size of the fiscal multiplier at the ZLB, it is crucial to understand the inflation response to government spending—and this inflation response is quite different for durable and nondurable goods. As can be seen in Panel B, government spending on nondurable goods raises inflation substantially more than spending on durable goods. Additionally, the impulse response for nondurable goods displays a slightly hump-shaped pattern with initial inflation and subsequent mild deflation as the government reduces spending. When good X is durable, inflation continues to rise for much longer, reflecting the fact that households rebuild their durables stock as soon as government demand falls.

For the fiscal multiplier to be large, inflation must be high when the interest rate is fixed at the ZLB. Conversely, deflation at the ZLB has strong contractionary effects. Panels C and D show the multiplier and the price path when the ZLB binds for 8 quarters. The multiplier for the moderately durable good with a service life of 10 years is now above one and the nondurables multiplier is near 3.

[22]More precisely, the impulse response functions are computed as the difference between the response with the fiscal policy shock and that without the fiscal policy shock. The resulting price dynamics are therefore entirely caused by the spending and not the discount factor shock.

The multiplier for nondurables falls when the ZLB binds for 16 quarters (Panel E). The reason is that prices are falling while the economy is at the ZLB. In fact, when the ZLB binds for 16 quarters, the multiplier for nondurables falls below that for moderately durable goods with $\delta_D = 0.025$. The catch-up of private sector spending prevents inflation from falling as much as it does when the government purchases nondurable goods. For the long-lived durable the multiplier remains below one.

In summary, the multiplier remains relatively small at the ZLB when the good in question is sufficiently durable. Yet, for goods with intermediate durability the multiplier can be above one if the ZLB binds sufficiently long.

1.5 Conclusion

Both neoclassical and New Keynesian models predict that the fiscal multiplier for temporary increases in spending is much smaller when the government buys durable rather than nondurable goods. I show that empirical evidence confirms this prediction. In U.S. aggregate data the fiscal multiplier is about 0.5 smaller if the government purchases durable rather than nondurable goods. At the industry level, spending in durables industries also leads to substantially smaller increases in economic activity than spending in nondurables industries.

These results raise significant concerns about the effectiveness of fiscal stimulus that is targeted towards infrastructure. As many other stimulus programs, the American Recovery and Reinvestment Act contained provisions to raise spending on highly durable goods such as highway infrastructure, high-speed rail corridors, railroads, airports, and broadband.[23] More generally, it has been suggested to assemble a pool of "shovel-ready projects" to be implemented when the economy next plunges into recession. The findings in this paper

[23] This information is taken from `http://www.recovery.gov/arra/Transparency/fundingoverview/Pages/fundingbreakdown.aspx`. Since the website is not operated anymore, I accessed a cached version (from 01/04/2014) via `https://archive.org/`.

suggest that such policies are unlikely to have substantial effects on aggregate demand.

This paper also provides guidance for future research on fiscal policy. The large difference in the sizes of durables and nondurables multipliers imply that future work should distinguish between these types of spending. Since the composition of spending matters, estimates of multipliers for total spending suffer from an external validity problem. If the composition of spending changes, the multiplier changes as well.

Finally, it is likely that other product dimensions matter. As a general rule, industries with more elastic supply curves and less elastic demand curves should have larger fiscal multipliers. Price stickiness, factor mobility, and tradability are just three properties likely to be associated with the elasticities of demand or supply.

Table 1.1: Baseline calibration

Parameter	Value	Description
β	0.99	Discount factor
η	1	Labor supply elasticity
σ	0.25	Intertemporal elasticity of substitution
ρ	0.9999	Elasticity of substitution between goods in the X and the Z sector
α	1/3	Capital share
χ	0.45	$1-\chi$ is the cost share of intermediates
ε	6	Elasticity of substitution in aggregator
θ_X, θ_Z	2/3	Price stickiness
δ_K	0.025	Depreciation rate of capital
X_H/X	0.80	Fraction of private sector spending in steady state
ω	0.9693	Preference weight on consumption of Z sector goods
ϱ_X	0.75	Persistence of fiscal policy shock
ϕ_π	1.1	Taylor rule response coefficient on inflation
ϕ_Y	0	Taylor rule response coefficient on output
ζ_K	20	Capital adjustment costs

Table 1.2: Estimates of durable and nondurable goods multipliers

	Sample					
	1929-2014			1947-2014		
Dependent variable	μ_X	μ_C	ν	μ_X	μ_C	ν
Output	0.04	0.49	0.065	-1.19	0.83	0.041
	(0.32)	(0.08)	(0.011)	(2.24)	(0.43)	(0.007)
Private consumption	-0.63	0.02	0.024	-1.13	0.14	0.012
	(0.20)	(0.04)	(0.008)	(1.03)	(0.20)	(0.003)
Private investment	-0.89	0.00	0.032	-1.12	0.09	0.028
	(0.26)	(0.05)	(0.004)	(0.93)	(0.18)	(0.003)
Private consumption of	-0.54	0.03	0.022	-0.95	0.17	0.009
nondurables and services	(0.16)	(0.04)	(0.006)	(0.95)	(0.19)	(0.003)
Private investment plus	-0.98	0.00	0.034	-1.30	0.07	0.030
private durables consumption	(0.33)	(0.06)	(0.005)	(1.01)	(0.20)	(0.003)
Net exports	0.00	-0.05	-0.002	-0.23	0.01	-0.007
	(0.08)	(0.01)	(0.003)	(0.32)	(0.08)	(0.001)

Notes: The table reports the coefficient estimates of μ_X, μ_C, and ν as defined in equation (1.19). Newey-West standard errors are reported in parentheses.

Table 1.3: First stages

First stages for nondurable goods industries

First stage dep. variable		Horizon h	0	1	2	3	4
$\frac{G_{i,t}-G^T_{i,t-1}}{VA^T_{i,t-1}}$	AP F-statistic		21.3	30.5	38.6	46.3	39.2
	p-value		0.00	0.00	0.00	0.00	0.00
	R-squared		0.42	0.43	0.44	0.44	0.44
$\frac{G_{i,t-1}-G^T_{i,t-2}}{VA^T_{i,t-2}}$	AP F-statistic		15.0	17.6	22.5	21.9	22.0
	p-value		0.00	0.00	0.00	0.00	0.00
	R-squared		0.42	0.42	0.42	0.43	0.43
$\frac{G_{i,t-2}-G^T_{i,t-3}}{VA^T_{i,t-3}}$	AP F-statistic		41.1	54.8	54.8	54.5	62.9
	p-value		0.00	0.00	0.00	0.00	0.00
	R-squared		0.38	0.43	0.43	0.44	0.44
Observations			967	932	897	862	827

First stages for durable goods industries

First stage dep. variable		Horizon h	0	1	2	3	4
$\frac{G_{i,t}-G^T_{i,t-1}}{VA^T_{i,t-1}}$	AP F-statistic		225.5	249.3	353.9	402.8	271.2
	p-value		0.00	0.00	0.00	0.00	0.00
	R-squared		0.44	0.48	0.52	0.54	0.54
$\frac{G_{i,t-1}-G^T_{i,t-2}}{VA^T_{i,t-2}}$	AP F-statistic		225.8	280.4	282.8	229.2	162.8
	p-value		0.00	0.00	0.00	0.00	0.00
	R-squared		0.48	0.51	0.53	0.53	0.55
$\frac{G_{i,t-2}-G^T_{i,t-3}}{VA^T_{i,t-3}}$	AP F-statistic		339.4	375.0	456.4	349.6	342.2
	p-value		0.00	0.00	0.00	0.00	0.00
	R-squared		0.52	0.53	0.54	0.56	0.56
Observations			2115	2039	1963	1887	1811

Notes: The table shows the first stages of the 2SLS estimator of specification (1.20) when the dependent variable is value added. For all other dependent variables, the F-statistics are virtually identical. The instruments are (1.21) and its two lags. AP F-statistic stands for Angrist-Pischke F-statistic of excluded instruments and the subsequent lines show the associated p-values.

Table 1.4: Industry-level fiscal multipliers

Durable goods multipliers			
Years after shock	1	2	3
Gross output	0.25	0.24	0.51
	$[-0.56, 1.14]$	$[-0.46, 1.18]$	$[-0.22, 1.52]$
Value added	0.35	0.36	0.57
	$[-0.07, 0.75]$	$[0.00, 0.77]$	$[0.19, 1.02]$
Cost of materials	-0.01	-0.06	-0.03
	$[-0.54, 0.68]$	$[-0.52, 0.66]$	$[-0.45, 0.76]$
Energy expenditures	0.007	0.005	0.007
	$[-0.005, 0.013]$	$[-0.004, 0.011]$	$[-0.002, 0.014]$
Employment	1.76	3.17	5.83
(employees per year per $1m)	$[-3.86, 4.24]$	$[-1.24, 6.19]$	$[1.51, 9.29]$

Nondurable goods multipliers			
Years after shock	1	2	3
Gross output	2.63	2.72	3.21
	$[0.47, 5.47]$	$[0.12, 5.45]$	$[0.46, 6.60]$
Value added	0.99	0.96	1.28
	$[0.05, 2.58]$	$[-0.27, 2.33]$	$[0.00, 2.88]$
Cost of materials	1.64	1.57	1.90
	$[-0.58, 4.68]$	$[-1.04, 4.51]$	$[-0.88, 5.54]$
Energy expenditures	0.056	0.060	0.047
	$[-0.014, 0.114]$	$[-0.011, 0.140]$	$[-0.021, 0.110]$
Employment	16.01	15.87	18.22
(employees per year per $1m)	$[5.26, 33.64]$	$[1.47, 31.21]$	$[3.19, 33.32]$

Notes: The table reports the cumulative multipliers for various outcome variables. Multipliers for gross output, value added, cost of materials, and energy expenditures have the usual interpretation of one additional dollar in the outcome variable per additional dollar of military spending. The multiplier for employment is expressed as the number of employees per year per $1 million of military spending. 80 percent confidence intervals are reported in square brackets. They are calculated using a blocks-of-blocks bootstrap (see *Berkowitz et al.* (1999) and the references cited therein) with 2000 bootstrap samples.

Table 1.5: Industry-level multipliers for durable goods in recessions and expansions

Recession multipliers			
Years after shock	1	2	3
Gross output	-2.82	-1.82	-0.87
	$[-8.04, 0.81]$	$[-6.42, 1.05]$	$[-3.83, 3.26]$
Value added	-1.10	-0.59	-0.08
	$[-3.15, 0.41]$	$[-2.65, 0.59]$	$[-2.26, 1.38]$
Cost of materials	-0.87	-0.32	0.15
	$[-3.17, 1.33]$	$[-2.36, 1.66]$	$[-2.36, 2.73]$
Energy expenditures	-0.009	-0.004	0.009
	$[-0.052, 0.027]$	$[-0.037, 0.032]$	$[-0.052, 0.053]$
Employment	5.33	6.70	10.53
(employees per year per $1m)	$[-10.38, 23.64]$	$[-6.16, 26.53]$	$[-16.82, 31.20]$
Expansion multipliers			
Years after shock	1	2	3
Gross output	1.79	1.15	0.82
	$[-0.60, 3.81]$	$[-0.93, 3.14]$	$[-1.36, 2.76]$
Value added	0.96	0.67	0.67
	$[-0.27, 2.02]$	$[-0.27, 1.74]$	$[-0.44, 1.67]$
Cost of materials	0.21	-0.32	-0.63
	$[-1.78, 1.79]$	$[-1.87, 1.34]$	$[-1.90, 1.20]$
Energy expenditures	0.019	0.011	0.007
	$[-0.016, 0.042]$	$[-0.015, 0.030]$	$[-0.020, 0.022]$
Employment	-1.63	0.10	3.93
(employees per year per $1m)	$[-17.75, 5.46]$	$[-13.70, 5.94]$	$[-9.25, 9.45]$

Notes: The table reports the cumulative multipliers for various outcome variables. Multipliers for gross output, value added, cost of materials, and energy expenditures have the usual interpretation of one additional dollar in the outcome variable per additional dollar of military spending. The multiplier for employment is expressed as the number of employees per year per $1 million of military spending. 80 percent confidence intervals are reported in square brackets. They are calculated using a blocks-of-blocks bootstrap (see *Berkowitz et al.* (1999) and the references cited therein) with 2000 bootstrap samples.

Figure 1.1: Crowding out of private sector spending.

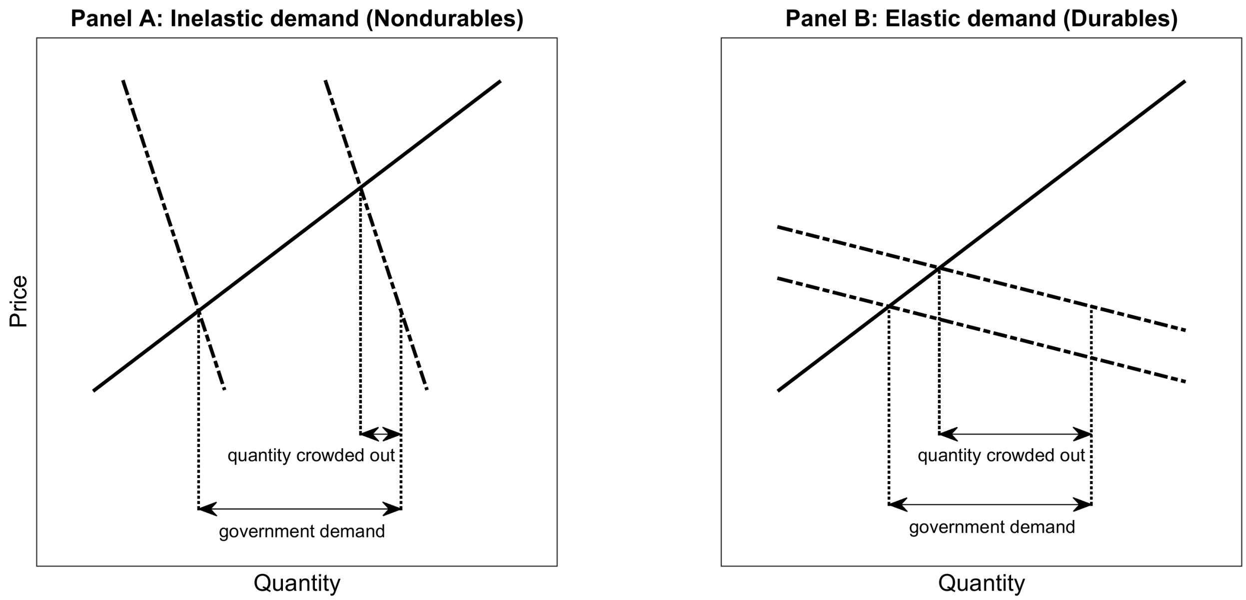
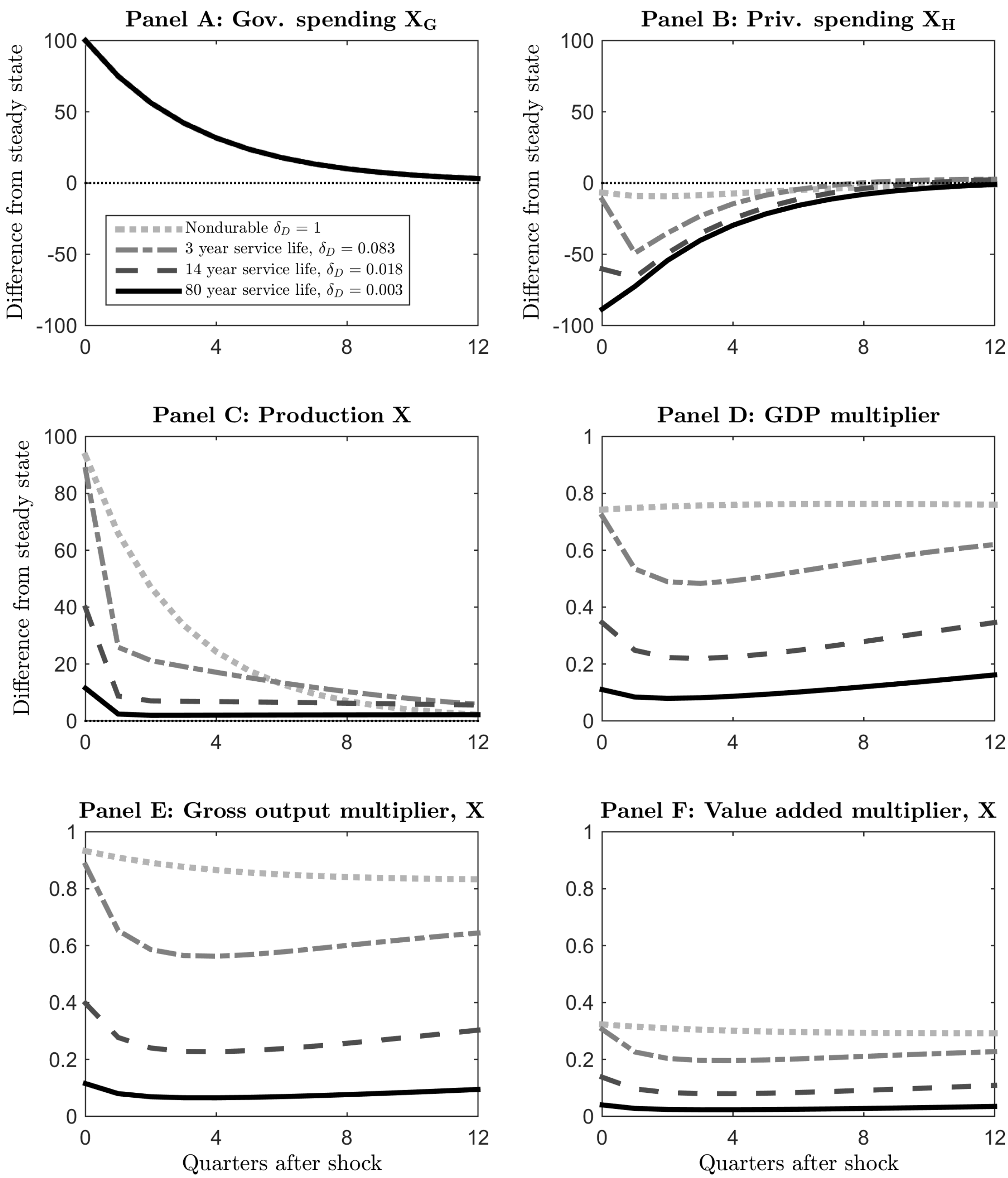
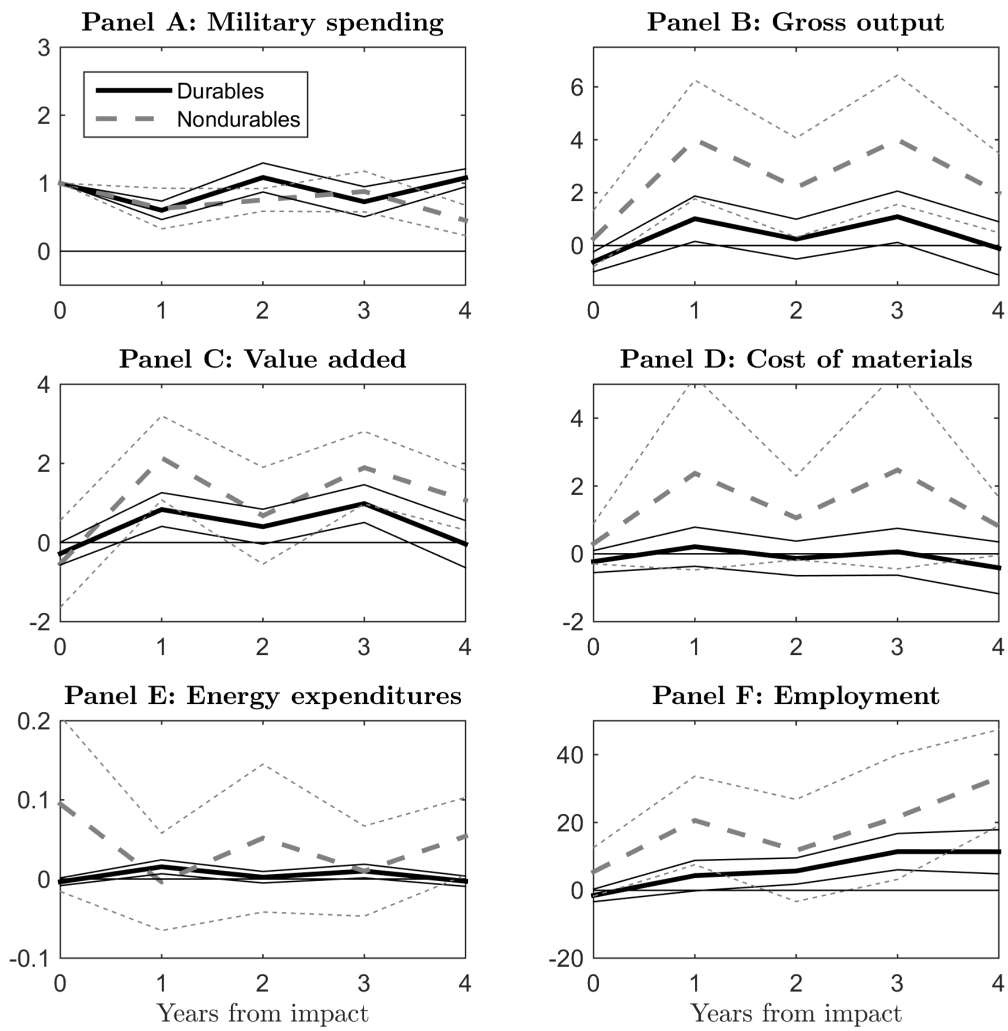


Figure 1.2: Impulse response functions for a government spending shock



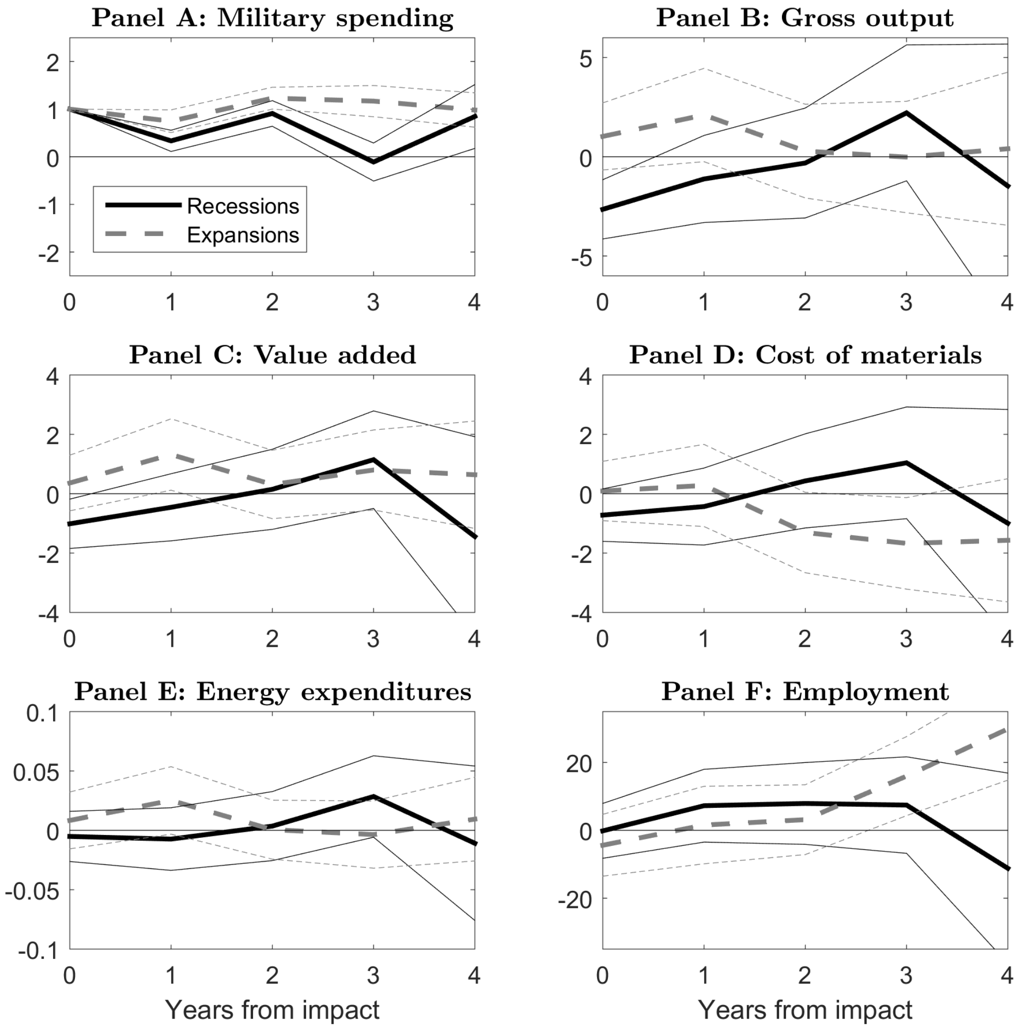
Notes: The figure plots impulse response functions for various calibrations. Panels A to C are expressed in absolute deviations from steady state. The impulse is a 100 unit increase of government spending in sector X.

Figure 1.3: Impulse response functions for durable and nondurable goods



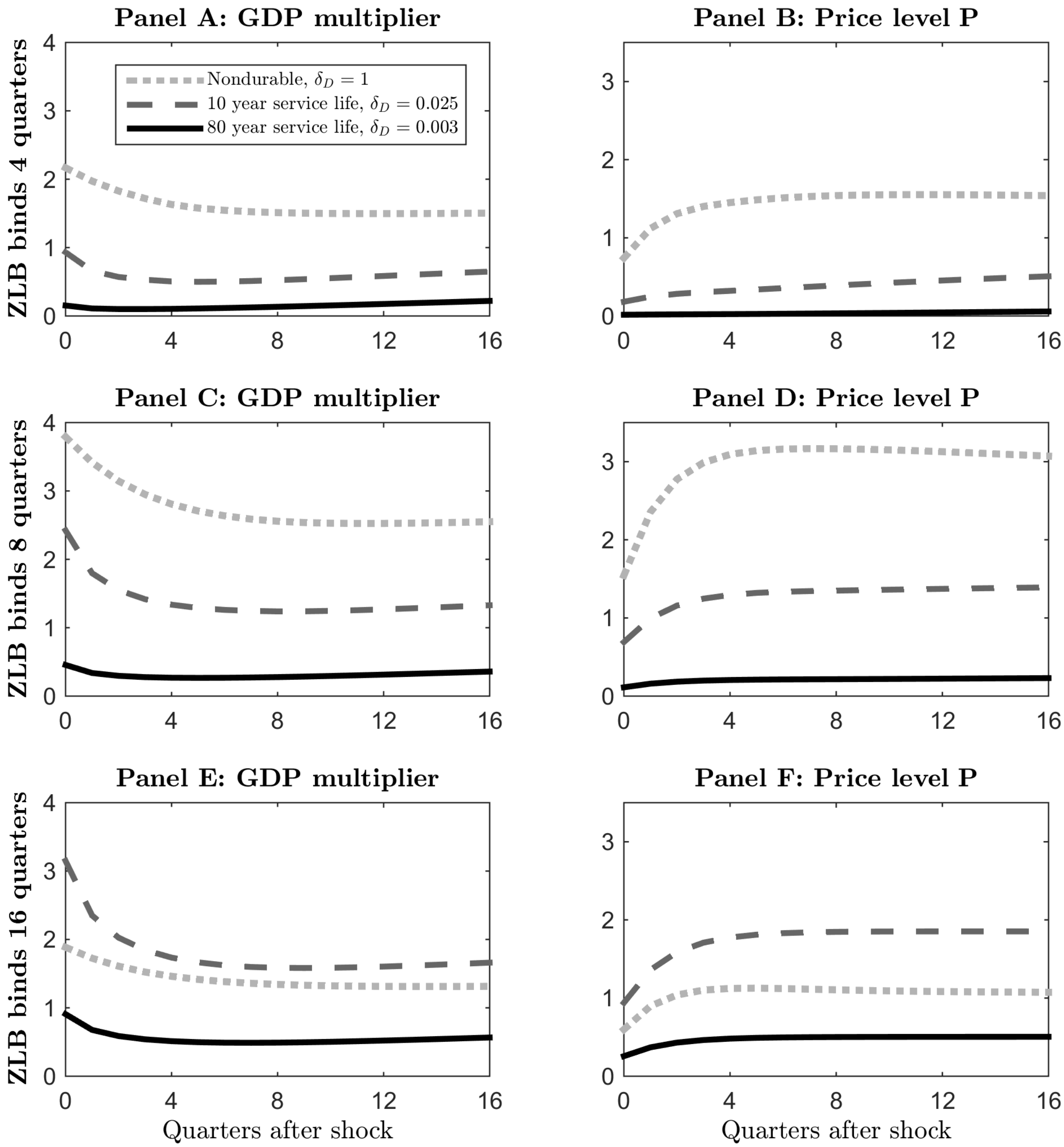
Notes: The figure plots impulse response functions for the baseline sample estimated using specification (1.20). See text for a description of the baseline sample. The shock is a unit increase of government spending above trend, normalized by the industry's value added. For production employment the impulse is a $1 million increase in military spending and the response is expressed as the number of additional employees. Shaded regions mark 80 percent confidence bands based on standard errors that are clustered at the industry level.

Figure 1.4: Impulse response functions for durable goods in recessions and expansions



Notes: The figure plots impulse response functions for the durable goods industries of the baseline sample estimated using specification (1.22). See Section 1.3 for a description of the baseline sample. The shock is a unit increase of government spending above trend, normalized by the sector's value added. For employment the impulse is a \$1 million increase in military spending and the response is expressed as the number of additional employees. Shaded regions mark 80 percent confidence bands based on standard errors that are clustered at the industry level.

Figure 1.5: Model impulse response functions



Notes: The figure fiscal multipliers and impulse response functions for various depreciation rates. Prior to time 0, the household's discount factor is first shocked so that the ZLB begins to bind. The persistence of the shock is chosen to imply that the ZLB binds for 4, 8, and 16 quarters beginning at time 0. At time 0 the government raises spending on X by 100 basis points. The impulse response functions for prices are expressed in relative deviations from steady state and the units are basis points.

CHAPTER II

Optimal Taylor Rules in New Keynesian Models

with Christopher L. House

2.1 Introduction

Taylor rules are simple linear relationships between a central bank's choice of a target interest rate, observed output (or the "output gap") and observed inflation (*Taylor* 1993). Suitably parameterized, they are reasonable descriptions of how actual central banks set interest rates. In addition, monetary policy is often discussed by journalists, researchers and central bankers themselves in terms that fit comfortably into a Taylor rule framework. For example, whether the federal funds rate should be increased or decreased is typically discussed in terms of whether inflation is relatively too high or whether GDP (or employment) is too low. A central banker who fights inflation aggressively can be modeled as following a Taylor rule with a large coefficient on inflation. A central banker who reacts more strongly to output would have a relatively higher coefficient on the output gap, and so forth.

While Taylor rules are useful descriptions of actual policy and common components of many prominent New Keynesian models, it is well-known that *optimal* monetary policy is rarely given by a Taylor rule. Instead, optimal policy depends in complicated ways on the underlying state variables and is often history dependent (see *Woodford* 1999). By confining attention to current inflation and the current output gap, a Taylor rule is unnecessarily restrictive.[1] Nevertheless, given the underlying model, it is a meaningful question to ask what the optimal parameterization of a Taylor rule is. Addressing this question is the goal of this paper. We anticipate that much of what we have to say will not come as a surprise to researchers at the frontier of New Keynesian economics. Indeed, we suspect that many of our results exist as folk wisdom among New Keynesian researchers. (If you are Mike Woodford or Jordi Galí, you can stop reading now). Instead, our intended audience consists of consumers of New Keynesian economics — researchers who often use macroeconomic models with sticky prices and want to draw on established results from the New Keynesian literature (we include ourselves in this group).

In this paper, we consider the nature of the optimal Taylor rule in the basic New Keynesian model. That is, we assume the monetary authority is committed to using a Taylor rule, and ask what coefficients maximize the central bank's objective function. When the output gap and inflation are observed without error, it is typically best to adopt infinitely aggressive responses to output and inflation i.e., the optimal coefficients on inflation and output are arbitrarily large. If the only shocks to the economy are shocks to the efficient rate of output ("demand shocks"), then any Taylor rule which responds infinitely strongly to either output or inflation (or both) will maximize welfare. If the only shocks perturbing the economy are shocks to the New Keynesian Phillips Curve ("cost-push shocks"), then there exists a continuum of optimal coefficients for the Taylor rule. If the output gap and inflation are observed with error, then the Taylor coefficients are finite. As the variance of measurement error grows, the optimal coefficients fall and the central bank reacts less to measured inflation and the output gap. We extend our analysis to the case in which the Taylor rule is expressed in terms of estimated output and estimated inflation. In this extension, the central bank solves

[1] See *Svensson* (2003) for a more general criticism of Taylor rules.

a signal extraction problem to estimate the output gap and inflation and sets the interest rate as a function of these estimates. Under such a policy rule, the optimal responses to estimated inflation and the estimated gap are again infinite. Because filtered estimates are typically revised gradually over time, a central bank that responds to estimates of inflation and output will often appear to be adhering to a form of interest rate smoothing. Indeed, there are special cases in which a Taylor rule with interest rate smoothing can duplicate the policy of responding to estimated inflation and the output gap.

Under an optimal Taylor rule, estimated deviations of output and inflation from their targets should be strongly negatively correlated. Intuitively, the optimal Taylor rule eliminates the effects of estimated demand shocks on inflation and the output gap. At the optimum, only variation due to estimated cost-push shocks should remain, implying that estimated inflation and output should move in opposite directions. Actual data on estimates of inflation and the gap are not strongly correlated, suggesting that the central bank is not reacting to demand shocks as aggressively as it should.

Our work is related to a large literature on optimal monetary policy and instrument rules. The contributions of *Giannoni and Woodford* (2003a,b) , *Woodford* (2003, Ch. 7) and *Giannoni* (2010, 2012) are closest to our work. Unlike us, *Giannoni and Woodford* (2003b) assume that the central bank has an explicit preference for smooth interest rates. Our paper highlights the role played by measurement error in tempering the central bank's optimal reaction to changes in the measured output gap and inflation. This type of measurement error has been termed data uncertainty by some researchers to distinguish it from other sources of policy uncertainty (such as parameter and model uncertainty, see *Dennis* (2005) and the references cited therein). *Orphanides* (2001, 2003) shows that real-time measures of inflation and the output gap are sufficiently noisy to justify relatively small Taylor rule coefficients. *Rudebusch* (2001), *Smets* (2002) and *Billi* (2012) all conclude that measurement error naturally encourages central banks to adopt less aggressive policy reaction rules.[2]

Our paper is also closely related to the literature on signal extraction and optimal monetary policy. *Aoki* (2003) considers optimal monetary policy with signal extraction in an environment similar to ours. There are several differences between Aoki's work and ours. First, unlike our model, Aoki's does not have shocks to the New Keynesian Phillips Curve and thus the so-called divine coincidence holds. If the central bank successfully stabilizes prices, it simultaneously eliminates the output gap. Second, while Aoki discusses optimal policy under discretion, we restrict attention to the optimal Taylor rule a form of commitment. Third, when formulating the signal extraction problem, Aoki assumes that the central bank learns the true values of output and inflation with a one period lag. In our formulation, these values are never revealed. Methodologically, we draw heavily on results in *Svensson and Woodford* (2003, 2004). Although they do not consider restricted instrument rules such as the Taylor rule, several of their findings continue to hold in our setting.

2.2 Baseline Model And Policy Objective

Our analysis is based on simple variations of the standard New Keynesian framework consisting of a Phillips Curve and the New Keynesian IS curve. These log-linear aggregate relationships are typically derived under the assumption that firms infrequently adjust prices according to the *Calvo* (1983) mechanism.[3] (See *Galí* (2008) or *Woodford* (2003) for standard treatments of this derivation.)

[2]Both *Rudebusch* (2001) and *Smets* (2002) numerically analyze optimal policy based on estimated New Keynesian systems. *Billi* (2012) numerically analyzes optimal monetary policy at the zero lower bound on nominal interest rates. Both Billi and Smets assume that the central bank wants to minimize interest rate variation in addition to variation in output and inflation. See *Taylor* (2001) for additional work on optimal monetary policy rules.

[3]The standard New Keynesian model abstracts from investment in physical capital and durable goods. While this assumption is common, it has important consequences for the analysis of the model and for optimal policy. See *Barsky et al.* (2003, 2007) and *Barsky et al.* (2014) for a more detailed discussion of the consequences of this assumption.

The Phillips curve relates inflation, π_t, to the output gap,y_t, expected future inflation, and a cost-push shock,u_t,

$$\pi_t = \kappa y_t + \beta \mathbb{E}_t[\pi_{t+1}] + u_t \tag{2.1}$$

where $\kappa > 0$ is the macroeconomic rate of price adjustment. The presence of the cost-push shock implies a trade-off between the stabilization of the output gap and inflation (e.g. *Gertler et al.* (1999)). *Woodford* (2003, Ch. 6) shows how the cost-push shock can be motivated from first principles.

The New Keynesian IS curve is given by

$$y_t = \mathbb{E}_t[y_{t+1}] - \frac{1}{\sigma}\left(i_t - \rho - r_t^e - \mathbb{E}_t[\pi_{t+1}]\right). \tag{2.2}$$

Here $\rho + r_t^e$ is the efficient rate of interest the interest rate consistent with the level of output that would prevail under perfect price flexibility in the absence of all other distortions. We express this rate as the sum of the rate of time preference ρ and a shock r_t^e which is centered at zero to simplify the exposition of later results. Below we often (somewhat imprecisely) refer to r_t^e as the efficient rate. The remaining terms are i_t, the nominal interest rate, and σ, the coefficient of relative risk aversion (equivalently, the inverse of the intertemporal elasticity of substitution).

The efficient rate shock and the cost-push shock are assumed to follow the AR(1) processes

$$r_{t+1}^e = \varrho_r r_t^e + \varepsilon_{t+1}^r, \qquad \varrho_r \in [0, 1), \tag{2.3}$$

$$u_{t+1} = \varrho_u u_t + \varepsilon_{t+1}^u. \qquad \varrho_u \in [0, 1). \tag{2.4}$$

We close the model by assuming that the monetary authority commits to a Taylor rule,

$$i_t = \rho + \phi_\pi \pi_t^m + \phi_y y_t^m. \tag{TR1}$$

Importantly, we distinguish between the actual output gap and the output gap observed by the monetary authority, and similarly between actual inflation and measured inflation. The central bank can respond only to measured output and inflation, so in (TR1) π_t^m and y_t^m denote *measured* inflation and the *measured* output gap. Using the analogy in *Bernanke* (2004), measurement error creates a foggy windshield through which the monetary authority sees the economy.

Orphanides (2001) suggests that modeling measurement error as additive is a reasonable approximation of reality. We follow Orphanides and assume that measured inflation and the measured output gap are $\pi_t^m = \pi_t + m_t^\pi$ and $y_t^m = y_t + m_t^y$, where m_t^π and m_t^y denote the respective measurement errors. Both types of measurement error follow AR(1) processes

$$m_{t+1}^\pi = \varrho_{m\pi} m_t^\pi + \varepsilon_{t+1}^{m\pi}, \qquad \varrho_{m\pi} \in [0,1)\,, \tag{2.5}$$

$$m_{t+1}^y = \varrho_{my} m_t^y + \varepsilon_{t+1}^{my}, \qquad \varrho_{my} \in [0,1)\,. \tag{2.6}$$

Unless stated otherwise, all error terms in the efficient rate, the cost-push shock and the measurement error processes are assumed to be uncorrelated.

A comprehensive literature has examined equilibrium determinacy in the New Keynesian model when monetary policy follows a Taylor rule. We will not dwell on this issue further and simply note that the well-known condition for a unique equilibrium is

$$\kappa\left(\phi_\pi - 1\right) + \left(1 - \beta\right)\phi_y > 0 \tag{2.7}$$

for the baseline model discussed in Section 4.[4] This condition is assumed to hold at all times unless otherwise stated.

[4]The argument in *Bullard and Mitra* (2002) continues to hold in our setting.

We assume that the central bank wishes to minimize an expected discounted sum of weighted squared inflation and the output gap

$$(1-\beta)\,\mathbb{E}\left[\sum_{t=0}^{\infty}\beta^{t}\left(\alpha\pi_t^2+y_t^2\right)\right]=\mathbb{E}\left[\alpha\pi_t^2+y_t^2\right]. \tag{2.8}$$

Here, α denotes the relative weight that the central bank places on inflation and $\mathbb{E}\,[.]$ is the unconditional expectations operator. An objective of this form can be derived as a quadratic approximation of the representative consumer's utility function (see *Rotemberg and Woodford* (1997) and *Woodford* (2003)). The optimal policy problem is then to choose ϕ_y and $\phi_p i$ to minimize (2.8) subject to (2.1), (2.2) and (TR1).

2.3 Equilibrium in the Baseline Model

Our framework is sufficiently simple to solve for the equilibrium analytically. Inflation and the output gap are linear functions of the four exogenous state variables r_t^e,u_t,m_t^{π}, andm_t^{y}. Applying the method of undetermined coefficients yields expressions for the behavior of inflation and the output gap. We characterize the model's solution in the following lemma. Proofs of all results are in the appendix.

Lemma 1: *Under the assumptions in Section 2.2, the unique competitive equilibrium of the model is given by the equations*

$$\begin{aligned}\pi_t =&\frac{\kappa}{\Phi_r+\phi_y\left(1-\beta\varrho_r\right)+\kappa\phi_\pi}r_t^e+\frac{\phi_y+\left(1-\varrho_u\right)\sigma}{\Phi_u+\phi_y\left(1-\beta\varrho_u\right)+\kappa\phi_\pi}u_t\\ &-\frac{\kappa\phi_\pi}{\Phi_{m\pi}+\phi_y\left(1-\beta\varrho_{m\pi}\right)+\kappa\phi_\pi}m_t^{\pi}-\frac{\kappa\phi_y}{\Phi_{my}+\phi_t\left(1-\beta\varrho_{my}\right)+\kappa\phi_\pi}m_t^{y},\end{aligned} \tag{2.9}$$

$$
\begin{aligned}
y_t = & \frac{1-\beta\varrho_r}{\Phi_r+\phi_y\left(1-\beta\varrho_r\right)+\kappa\phi_\pi} r_t^e + \frac{\varrho_u-\phi_\pi}{\Phi_u+\phi_y\left(1-\beta\varrho_u\right)+\kappa\phi_\pi} u_t \\
& - \frac{\left(1-\beta\varrho_{m\pi}\right)\phi_\pi}{\Phi_{m\pi}+\phi_y\left(1-\beta\varrho_{m\pi}\right)+\kappa\phi_\pi} m_t^\pi - \frac{\left(1-\beta\varrho_{my}\right)\phi_y}{\Phi_{my}+\phi_t\left(1-\beta\varrho_{my}\right)+\kappa\phi_\pi} m_t^y,
\end{aligned}
\tag{2.10}
$$

where $\Phi_j = \sigma\left(1-\varrho_j\right)\left(1-\beta\varrho_j\right) - \kappa\varrho_j, j \in \{r, u, m\pi, my\}$,*are constants that are independent of monetary policy.*

Examining the coefficients in (2.9) and (2.10) provides simple intuition about how monetary policy influences the economy. First, notice that the coefficients on the efficient rate shock (r_t^e) are positive in both (2.9) and (2.10) (all of the denominators in (2.9) and (2.10) are positive). Thus, an increase in the efficient rate (an aggregate demand shock) increases inflation and the output gap. Moreover, the Taylor rule coefficients ϕ_y and ϕ_π appear only in the denominators of the coefficients on the efficient rate shock, so strong responses by the central bank dampen fluctuations caused by shocks to the efficient rate.

Now consider the coefficients on the cost-push disturbances (u_t). Assuming that ϕ_y is positive and $\phi_\pi > \varrho_u$, cost-push shocks reduce output and raise inflation. Strong reactions to output (i.e., large values of ϕ_y) imply that the economy can enjoy reduced output gap volatility only at the expense of higher inflation volatility. Unlike shocks to the efficient rate of interest, cost-push shocks clearly entail a trade-off between inflation and output stabilization.

Finally, measurement error of either type impacts equilibrium inflation and the output gap negatively. The interpretation of this relationship is natural. For example, a positive innovation to m_π^t makes inflation appear higher than actual inflation. In response, the central bank raises interest rates causing both output and inflation to fall. Similar reasoning applies to measurement error in the output gap.

With Lemma 1 in hand, we now turn our attention to the central bank's optimal choices of ϕ_y and ϕ_π.

2.4 Optimal Taylor Rules in the Baseline Model

The optimal Taylor rule coefficients, ϕ_y^* and ϕ_π^*, minimize the central bank's objective (2.8) subject to the equilibrium conditions (2.9) and (2.10). To build intuition, we begin by considering four special cases. All of these special cases share the property that both inflation and the output gap are observed without error.

Special Case 1: Only shocks to the efficient rate of interest. Assume that the only shocks to the economy are shocks to the efficient rate of interest, r_t. In this case, equations (2.9) and (2.10) simplify to

$$\pi_t = \frac{\kappa}{\Phi_r + \phi_y (1 - \beta \varrho_r) + \kappa \phi_\pi} r_t^e,$$

$$y_t = \frac{1 - {}_R}{\Phi_r + \phi_y (1 - \beta \varrho_r) + \kappa \phi_\pi} r_t^e.$$

Clearly, it is optimal for the central bank to respond infinitely strongly to deviations in either inflation or the output gap (or both). This is the well-known divine coincidence case in which it is possible for the central bank to kill two birds with one stone by simply eliminating inflation variability (see e.g., *Blanchard and Galí* (2007)).

Special Case 2: Only Cost-Push Shocks. Optimal monetary policy is not so simple when there are shocks to the Phillips curve. Suppose the only shocks to the economy are cost-push shocks. The central bank can no longer eliminate both inflation and output variability and instead must choose whether to endure large swings in inflation to reduce the variation in output or vice versa. The best choice will depend on the underlying parameters of the model, particularly the relative weight α that the central bank attaches to inflation variability in its objective (2.8). In this case, (2.9) and (2.10) collapse to

$$\pi_t = \frac{\phi_y + (1-\varrho_u)\,\sigma}{\Phi_u + \phi_y\,(1-\beta\varrho_u) + \kappa\phi_\pi} u_t,$$

$$y_t = \frac{\varrho_u - \phi_\pi}{\Phi_u + \phi_y\,(1-\beta\varrho_u) + \kappa\phi_\pi} u_t.$$

The optimal Taylor coefficients are given in Proposition 1.

Proposition 1:

(i) If the only shocks to the model are cost-push shocks, then the optimal policy requires that the Taylor rule coefficients ϕ_y^* and ϕ_π^* lie on the affine manifold

$$\phi_\pi^* = \varrho_u + \frac{\alpha\kappa\sigma\,(1-\varrho_u)}{1-\beta\varrho_u} + \frac{\alpha\kappa}{1-\beta\varrho_u}\phi_y^*. \tag{2.11}$$

(ii) For any ϕ_y^* and ϕ_π^* satisfying (2.11), the equilibrium satisfies

$$\pi_t = \frac{1-\beta\varrho_u}{\alpha\kappa^2 + (1-\beta\varrho_u)^2} u_t, \qquad y_t = -\frac{\alpha\kappa}{\alpha\kappa^2 + (1-\beta\varrho_u)^2} u_t.$$

Any combination of coefficients on line (2.11) is equally desirable. Moreover, the equilibrium paths for output and inflation do not depend on the exact coefficient values, provided that they are on the manifold.

To understand the intuition for this result, consider a positive innovation to u_t. By itself, the cost-push shock puts upward pressure on inflation and downward pressure on output. The central bank then faces a dilemma: any interest rate response which closes one of the gaps, widens the other. If the bank cuts the interest rate to raise output, inflation rises. If the bank fights inflation instead, it must endure even lower production in the short term.

Suppose the bank decides to raise interest rates. Suppose further that it has chosen a pair of coefficients (ϕ_y^*, ϕ_π^*) which satisfy (2.11) and imply the desired increase in interest rates. According to Proposition 1, the same interest rate response can be achieved by many different Taylor rules. If the central bank instead opts for a stronger response to inflation (a higher ϕ_π) the nominal interest rate will increase a bit more. To re-establish the prior interest rate change, the central bank can adopt a stronger response to output. The higher ϕ_y implies a lower interest rate because cost-push shocks lower output whenever they raise inflation. Equation (2.11) simply gives all pairs of coefficients that result in the same interest rate.

Special Case 3: Shocks to both the IS curve and the Phillips curve. We next combine the previous two cases and consider (uncorrelated) shocks to both the IS-Curve (r_t^e) and the Phillips Curve (u_t). In Special Case 1, optimal policy required infinitely large Taylor Rule coefficients. In Special Case 2, optimal policy required coefficients on the upward sloping line (2.11). Not surprisingly, the optimal Taylor rule in this third case calls for arbitrarily large coefficients which also satisfy (2.11). Equation (2.11) implies that the optimal ratio of ϕ_π to ϕ_y converges to $\alpha\kappa/\left(1-\beta\varrho_u\right)$. This ratio is higher if the central bank places greater weight on inflation or if the tradeoff between inflation and output (captured by the parameter κ) is more favorable.

This policy will imply the same equilibrium paths as in Proposition 1: all of the disturbances originating from variation in the efficient rate will be eliminated and only the cost-push shocks affect output and inflation. Combining the two equations in Proposition 1 part (ii) yields

$$\pi_t = -\frac{1-\beta\varrho_u}{\alpha\kappa} y_t.$$

Hence, for the optimal Taylor rule, the observed output gap and inflation are perfectly negatively correlated.[5]

[5] Note that under the optimal policy, estimation of a New Keynesian Phillips curve will be particularly problematic. Typically, the structural shock u_t is correlated with the regressors y_t and $\mathbb{E}_t[\pi_{t+1}]$ but under optimal policy, both regressors are functions only of u_t and are therefore perfectly correlated with the error. The optimal Taylor rule eliminates all variation other than variation associated with u_t making the bias

Special Case 4: Correlated shocks in the i.i.d. case. For Special Case 3 we assumed uncorrelated cost-push and demand shocks. Here we briefly consider the optimal Taylor rule coefficients for transitory but correlated shocks. This case is relevant for later results and also shows that properties of the optimal Taylor rule can be extended to allow for correlated structural shocks.

Proposition 2: *Suppose r_t^e and u_t are i.i.d. over time and have covariance* $\mathrm{Cov}[r_t^e, u_t]$. *The Taylor rule coefficients satisfying*

$$\phi_\pi^* = \alpha\kappa\left(\phi_y^* + \sigma\right) + \left(1 + \alpha\kappa^2\right)\frac{\mathrm{Cov}[r_t^e, u_t]}{\mathrm{V}[u_t]} \tag{2.11'}$$

and $\phi_y^* \to \infty$ are optimal.

As before, the central bank must respond infinitely strongly to inflation and the output gap. Additionally, the relationship between ϕ_π^* and ϕ_y^* barely changes. To see this, notice that (2.11') simplifies to $\phi_\pi^* = \alpha\kappa\left(\phi_y^* + \sigma\right)$ when ϱ_u is set to zero (the first term on the right hand side of (2.11')). Hence, shock correlation only introduces the term $(1 + \alpha\kappa^2)\,\mathrm{Cov}[r_t^e, u_t]/\mathrm{V}[u_t]$ that was not present in the earlier condition. It is easy to verify that equilibrium outcomes only depend on the limit of ϕ_π^*/ϕ_y^* which is independent of the correlation of shocks. As in special cases 2 and 3, the output gap and inflation are perfectly negatively correlated in equilibrium.[6]

2.4.1 Monetary Policy in the Presence of Measurement Error

The simple New Keynesian framework captures many realistic features of monetary policy. The model embodies a tradeoff between inflation and output and suggests that the central bank has a particular advantage in minimizing economic instabilities that arise from

particularly pronounced.

[6]Even though ϕ_y and ϕ_π are infinite, the equilibrium interest rate will be finite. This can easily be seen easily in the i.i.d. case using (TR1), (2.9), (2.10) and (2.11').

demand shocks (shocks to the IS-Curve). Despite these attractive features, the model does not entail any costs to excessively strong reactions on the part of the central bank. In stark contrast to the modest empirical estimates of actual Taylor rules (see *Judd and Rudebusch* (1998), and more recently *Hofmann and Bogdanova* (2012)), optimal Taylor rule coefficients are often infinite.

In this section, we consider the model in which the output gap and inflation are measured with error. Measurement error is a natural candidate for why central banks do not respond more to observed changes in GDP and inflation. This concern was emphasized by *Friedman* (1953) who pointed out that activist policies might be destabilizing if policy actions were not sufficiently correlated with the true targets of policy. From observing equations (2.9) and (2.10) it is clear that, for any fixed coefficients ϕ_y and ϕ_π, greater measurement error reduces the correlation between the policy instrument and the targets and thus entails greater unwanted variation in output and inflation. Indeed, it would seem that if measurement error was sufficiently high, it would be optimal not to respond to observed variations in inflation and output at all.[7]

When we allow for arbitrary variation in the efficient rate, the cost-push shock, and both types of measurement error, an analytical solution of the optimal policy problem is generally not feasible and so we instead use numerical methods to characterize the optimal Taylor rule. To build intuition, however, we first consider another special case.

Proposition 3: *Suppose all shocks are white noise, that is,* $\varrho_r = \varrho_u = \varrho_{m\pi} = \varrho_{my} = 0$. *Then the minimization of (2.8) subject to (2.9) and (2.10) yields the following optimal Taylor*

[7] This might present a problem for equilibrium determinacy. It is well known that determinacy requires that the central bank responds sufficiently strongly to inflation and output. If measurement error is large however, the Fed might be caught between a rock and a hard place. It would be forced to choose between a locally indeterminate equilibrium on the one hand and destabilization resulting from reactions to erroneous signals on the other. Accurate measurement and state estimation are therefore at the heart of monetary stabilization policy. We consider optimal state estimation on the part of the central bank below.

rule coefficients

$$\phi_y^* = \frac{1}{\sigma} \frac{\mathrm{V}[r_t^e]}{\mathrm{V}[m_t^y]}, \tag{2.12}$$

$$\phi_\pi^* = \frac{\kappa}{\sigma} \frac{(\alpha\kappa^2 + 1)\,\mathrm{V}[m_t^y] + \alpha \mathrm{V}[u_t]}{(\alpha\kappa^2 + 1)\,\mathrm{V}[m_t^\pi] + \mathrm{V}[u_t]} \frac{\mathrm{V}[r_t^e]}{\mathrm{V}[m_t^y]} + \frac{\alpha\kappa\sigma \mathrm{V}[u_t]}{(\alpha\kappa^2 + 1)\,\mathrm{V}[m_t^\pi] + \mathrm{V}[u_t]}, \tag{2.13}$$

where $\mathrm{V}[\cdot]$ denotes the unconditional variance operator.

There are several striking features of the optimal Taylor rule in this setting. First, the optimal Taylor coefficients in this model are finite. The central bank avoids aggressive reactions to measured inflation and output because it knows that its actions would cause excessive fluctuations in actual output and inflation. As measurement error decreases, the Fed can adopt more and more aggressive reactions to inflation and output. It is also worth noting that only measurement error in the output gap is necessary for finite Taylor coefficients. Assuming that $\mathrm{V}[u_t] > 0$, measurement error in inflation is neither necessary nor sufficient for finite coefficients.

Second, the optimal choice of ϕ_y^* depends neither on α, nor on κ, nor on $\mathrm{V}[u_t]$. Instead, ϕ_y^* depends only on the ratio of the variance of shocks to the efficient rate, $\mathrm{V}[r_t^e]$, to the variance of measurement error in the output gap, $\mathrm{V}[m_t^y]$, together with the coefficient of relative risk aversion, σ. The reader might find the result in (2.12) somewhat counterintuitive. If risk aversion is relatively high then the household will strongly dislike output variability and presumably prefer a stronger output reaction. In contrast, the actual best reaction is decreasing in σ. The reason for this apparent contradiction is that we chose to specify the IS shocks as shocks to the efficient rate of interest itself (as is common in the literature). If we instead stated the shocks in terms of exogenous changes in the efficient growth rate of output, we would have $r_t^e = \sigma \mathbb{E}_t[\Delta y_{t+1}^e] = -\sigma y_t^e$ (the second equality uses the assumption that the autocorrelation of shocks is zero). In this case, the central bank's choice of ϕ_y^* can be written as

$$\phi_y^* = \sigma \frac{\mathrm{V}[y_t^e]}{\mathrm{V}[m_t^y]}$$

which is increasing in σ.

The optimal response to measured inflation (equation (2.13)) is somewhat more complex. We start with its more intuitive properties. First, as with ϕ_y^*, a larger variance of the efficient rate $\mathrm{V}[r_t^e]$ implies a stronger response to inflation. Second, greater measurement error in inflation requires more attenuated responses. Third, one can show that $\partial\phi_\pi^*/\partial\alpha > 0$ for *any* choice of model parameters so a stronger preference for price stability always implies stronger reactions to measured inflation.

The relationship of ϕ_π^* with the remaining parameters is less clear. To see how this coefficient depends on the shock variances consider the following limiting cases. Suppose first that cost-push shocks are dominant that is, consider the behavior of ϕ_π^* as $\mathrm{V}[u_t] \to \infty$. In this case, the optimal reaction to inflation approaches

$$\phi_\pi^* = \frac{\kappa\alpha}{\sigma}\frac{\mathrm{V}[r_t^e]}{\mathrm{V}[m_t^y]} + \kappa\alpha\sigma = \kappa\alpha\sigma\left(\frac{\mathrm{V}[y_t^e] + \mathrm{V}[m_t^y]}{\mathrm{V}[m_t^y]}\right),$$

where we have again used the relationship $r_t^e = -y_t^e$. The inflation response is increasing in the signal-to-noise ratio for the output gap, the weight the central bank places on inflation stability, the slope of the IS curve and the macroeconomic rate of price adjustment. Notice also that for large $\mathrm{V}[u_t]$ neither ϕ_y^* nor ϕ_π^* depend on measurement error in inflation.

Alternatively, consider the opposite extreme suppose there are no cost-push shocks at all. In this case,

$$\phi_\pi^* = \frac{\kappa}{\sigma}\frac{\mathrm{V}[r_t^e]}{\mathrm{V}[m_t^y]} = \kappa\sigma\frac{\mathrm{V}[y_t^e]}{\mathrm{V}[m_t^y]}.$$

Analogous to equation (2.12), the ratio of the variance of the efficient rate of output to that of measured inflation governs the strength of the reaction. In particular, as $\mathrm{V}[u_t]$ approaches zero, ϕ_π^* becomes independent of measurement error in the output gap. Finally, a larger macroeconomic rate of price adjustment, κ, raises the policy response.

Much of the optimal monetary policy literature assumes a quadratic objective function and linear constraints. In these settings the globally optimal policy exhibits certainty equivalence. That is, the presence and nature of additive stochastic disturbances does not affect optimal policy (see, e.g. *Ljungqvist and Sargent* (2004)). However, when optimal policy is restricted to a Taylor rule of the form (TR1), equations (2.9) and (2.10) show that the constraints are no longer linear in the choice variables. Hence, it is not surprising that certainty equivalence breaks down in our setting. This result is consistent with earlier findings (see, e.g. *Smets* (2002)).

2.4.2 Quantitative Analysis

We now analyze a calibrated version of the model to illustrate how the optimal Taylor rule changes with plausible variations in the model parameters. The length of one time period is a quarter. We assume logarithmic utility ($\sigma = 1$). The discount factor β is set to 0.99 and we set the macroeconomic rate of price adjustment to $\kappa = 0.34$.[8] We choose an autoregressive parameter of the efficient rate ϱ_r equal to 0.9. This value is similar to calibrations of trend stationary productivity shocks used in the real business cycle literature. Existing literature provides less guidance for the persistence of the cost-push shocks ϱ_u. Admitting that this choice is somewhat ad hoc we select a value of 0.5 as in one of the cases considered in *Galí* (2008, Ch. 5). The variances of the innovations to the efficient rate and the cost-push shocks are chosen so that the annual unconditional variances of r_t^e and u_t equal unity.

Orphanides (2003) provides estimates for the measurement error processes. Based on his calculations, measurement error in inflation is best approximated by a white noise process

[8]This value of can be derived from a Calvo model with a probability of price rigidity of 2/3 per quarter, together with our calibrated value of β, a Frisch labor supply elasticity of 1.00 and a linear production function.

with a quarterly standard deviation of roughly 0.50. In contrast, measurement error in the output gap is highly persistent with a quarterly autoregressive coefficient of about 0.95 and an innovation standard deviation of 0.66 (also quarterly). Finally, we assume that the monetary authority dislikes inflation and the output gap equally so $\alpha = 1$. Table 2.1 summarizes the baseline calibration.

The optimal policy coefficients, ϕ_π^* and ϕ_y^*, maximize objective (2.8) subject to the Phillips curve (2.1), the dynamic IS equation (2.2), and the Taylor rule (TR1). For the numerical solutions, we additionally impose condition (2.7) to ensure that the equilibrium is locally determinate. (This constraint never binds for the parameter values considered below.) For the benchmark calibration, the central bank's optimal Taylor rule has coefficients $\phi_\pi^* = 2.00$ and $\phi_y^* = 0.61$.[9]

We now examine how optimal policy depends on the model parameters. Figure 2.1 shows the optimal Taylor Rule coefficients as we vary the persistences of the shock processes in the neighborhood of the baseline calibration. To isolate the effect of shock persistence on policy, we adjust the variance of the innovations to ensure that the unconditional variance of the shocks is unchanged. In Figures 2.1 to 2.3, baseline parameter values are indicated by dotted vertical lines. All other parameters are held fixed at the level of the baseline calibration.

As the persistence of the efficient rate increases, the optimal reactions to inflation and the output gap both rise. In line with *Rudebusch* (2001), we interpret this finding as justifying intervention in the case of more persistent shocks while less intervention is necessary if the economy automatically and quickly reverts to its efficient allocation. However, this intuition does not always hold. For instance, the persistence of cost-push shocks has little influence on the optimal Taylor rule coefficients. Interestingly, persistent measurement error in inflation increases the optimal Taylor rule coefficients while persistent measurement error

[9] Whenever we report numerical values for the coefficient on the output gap, we annualize it by multiplying the quarterly value by four.

in the output gap reduces the optimal response to the output gap (and leaves the inflation coefficient almost unchanged).

Figure 2.2 shows how the optimal Taylor rule depends on the standard deviation of shock innovations. The upper left panel shows the Taylor rule coefficients as we change the volatility of the IS shocks. Not surprisingly (and incidentally, consistent with Proposition 3), a larger standard deviation of the efficient rate shocks raises ϕ_π^* and ϕ_y^*. The upper right panel considers changes in the standard deviation of cost-push shocks. Given our baseline calibration, as the cost-push shocks become more volatile the reaction to the output gap increases while the reaction to inflation falls. Intuitively, as the variance of cost-push shocks increases, the model behaves more and more like the special case in which there are only cost-push shocks. In that case, the optimal coefficients were restricted to the manifold (2.11). At the same time, the logic of Proposition 3 suggests finite Taylor rule coefficients as long as the output gap is observed with error. A combination of these two arguments seems to imply that as the variance of cost-push shocks becomes greater, the coefficients will approach a specific finite point on the line (2.11).

The bottom panels in Figure 2.2 show how the optimal Taylor rule depends on measurement error. Not surprisingly, for both types of measurement error, greater data uncertainty implies less active policy. It is worth noting that, at least for our baseline calibration, measurement error in the output gap is much more influential than measurement error in inflation. This is again consistent with the analytical result in Proposition 3.

Figure 2.3 depicts the optimal Taylor rule for alternative values of the weight on inflation in the objective function (α) and the macroeconomic rate of price adjustment (κ). The left panel illustrates the coefficients when we change the weight on inflation. As the preference for price stability increases (a greater α), the central bank adopts a Taylor rule with a higher reaction coefficient on inflation and a lower reaction coefficient on the output gap. The panel on the right shows how the coefficients vary with the macroeconomic rate of price adjustment. While the relationship is not monotone over the range considered in the figure, for relatively high values of κ, the central bank again chooses a stronger reaction to inflation and a weaker reaction to the output gap.

2.5 Signal Extraction and Optimal Taylor Rules

To this point, we have assumed that the central bank directly responds to *measured* inflation π_t^m and the *measured* gap y_t^m. Alternatively, we could consider a modified Taylor rule which stipulates that the central bank sets the interest rate as a function of *estimated* inflation and the output gap. Among others, *Orphanides* (2001) has advocated this specification. Drawing on the results of *Svensson and Woodford* (2003, 2004), we consider signal extraction and optimal Taylor rules in this section.[10]

In our model, agents in the private sector have full information. Their information set includes all variables dated t or earlier and all model parameters. In contrast, the central bank observes only measured inflation and the measured output gap. We assume that the central bank uses the Kalman filter to construct estimates of the true values of the output gap and the inflation rate. For a generic variable x_t, we let $x_{t|t}$ denote the central bank's estimate of the variable given all of the information available at date t.[11] Having solved the signal extraction problem, the central bank sets the interest rate according to the modified Taylor rule

$$i_t = \rho + \psi_\pi \pi_{t|t} + \psi_y y_{t|t}. \tag{TR2}$$

ψ_π and ψ_y are the Taylor rule coefficients which operate on the estimates $\pi_{t|t}$ and $y_{t|t}$. We list this Taylor rule as (TR2) to distinguish it from the more conventional Taylor rule (TR1).

[10] A number of researchers have examined signal extraction problems of central banks. Further references include *Swanson* (2000), *Aoki* (2003), and *Smets* (2002).

[11] Formally, the central bank's information set is $I_t^{CB} = \{\Theta, \pi_{t-j}^m, y_{t-j}^m : j \geq 0\}$ where Θ is a vector of all model parameters. Then, for any variable x_t, the central bank's date t estimate is $x_{t|t} = \mathbb{E}[x_t | I_t^{CB}]$. The observation equations are $\pi_t^m = \pi_t + m_t^\pi$ and $y_t^m = y_t + m_t^y$.

The remaining model equations are unchanged.

2.5.1 Optimal Policy

We begin by characterizing the central bank's estimates of inflation and the output gap as a function of the estimated disturbances. Lemma 2 is analogous to Lemma 1 for the model without noisy observations.

Lemma 2: *The central bank's estimates of inflation and the output gap satisfy*

$$\pi_{t|t} = \frac{\kappa}{\Phi_r + \psi_y (1 - \beta \varrho_r) + \kappa \psi_\pi} r^e_{t|t} + \frac{\psi_y + (1 - \varrho_u)\, \sigma}{\Phi_u + \psi_y (1 - \beta \varrho_u) + \kappa \psi_\pi} u_{t|t}, \tag{2.14}$$

$$y_{t|t} = \frac{1 - \beta \varrho_r}{\Phi_r + \psi_y (1 - \beta \varrho_r) + \kappa \psi_\pi} r^e_{t|t} + \frac{\varrho_u - \psi_\pi}{\Phi_u + \psi_y (1 - \beta \varrho_u) + \kappa \psi_\pi} u_{t|t}, \tag{2.15}$$

where Φ_r and Φ_u are defined as in Lemma 1.

The lemma shows that, with a suitable reinterpretation of the shocks, the equilibrium paths of the filtered variables $y_{t|t}$ and $\pi_{t|t}$ obey the same equilibrium conditions as the actual underlying variables y_t and π_t (see equations (2.9) and (2.10)).

We are now interested in finding the coefficients $\{\psi_y^*, \psi_\pi^*\}$ that minimize (2.8) subject to (2.1) to (2.6), (TR2), and the central bank's informational constraints. The following proposition presents a result for the i.i.d. case in which we can obtain a closed-form solution.

Proposition 4: *Suppose all shocks are contemporaneously uncorrelated with each other and i.i.d over time. Then the coefficients $\{\psi_y^*, \psi_\pi^*\}$ satisfying*

$$\psi_\pi^* = \alpha\kappa \left(\psi_y^* + \sigma\right) + \left(1 + \alpha\kappa^2\right) \frac{\mathrm{Cov}[r^e_{t|t}, u_{t|t}]}{\mathrm{V}\left[u_{t|t}\right]} \tag{11''}$$

and $\psi_y^* \to \infty$ are optimal.

Proposition 4 shows that, there is an optimal Taylor rule in terms of filtered output $y_{t|t}$

and inflation $\pi_{t|t}$ which embodies the same properties as Proposition 2 (Special Case 4). As in Special Case 4, the optimal Taylor rule coefficients are again infinitely large and ψ_π^*/ψ_y^* converges to $\alpha\kappa$. Notice the following subtle difference: in Proposition 2, we assumed an exogenous correlation between r_t^e and u_t. In contrast, the correlation between the estimates $r_{t|t}^e$ and $u_{t|t}$ in Proposition 4 arises endogenously even though the underlying shocks r_t^e and u_t are uncorrelated.

The correlation between $r_{t|t}^e$ and $u_{t|t}$ comes from the central bank's effort to infer the true shocks. For example, suppose the central bank observes positive inflation and a negative output gap. A negative supply shock ($u_t > 0$) is a natural candidate for such observations. Another possibility is the occurrence of a positive demand shock together with a negative innovation to measurement error in the output gap. Because the central bank attaches positive probability to many potential combinations of shocks, its estimates $r_{t|t}^e$ and $u_{t|t}$ will typically be correlated. For the optimal Taylor rule, the induced correlation is immaterial. Equations (2.14) and (2.15) with $\varrho_u = \varrho_r = 0$ imply that only the limit of ψ_π^*/ψ_y^* affects the central bank's estimates of inflation and the output gap.

We next turn to the general case in which all shocks have arbitrary autocorrelation. The analytical solution of the optimal policy problem is difficult so we use numerical methods to characterize the optimal Taylor rule.

Figure 2.4 shows level curves of the central bank's objective function when all parameters are set to the values of our baseline calibration summarized in Table 2.1. As we have seen in the i.i.d. case (Proposition 4), stronger responses to both expected inflation and the expected output gap reduce the central bank's loss function (2.8). This continues to be the case when there are persistent innovations. Again, the optimal Taylor rule coefficients are infinitely large.[12]

As the Taylor rule coefficients approach infinity, (2.14) and (2.15) imply that expected

[12]In all numerical cases we consider, the optimal Taylor rule coefficients approach infinity.

inflation and the expected output gap depend only on the estimated supply shock. In particular, $\pi_{t|t} = Au_{t|t}$ and $y_{t|t} = -Bu_{t|t}$ for some strictly positive constants A and B. Eliminating $u_{t|t}$ immediately yields

$$y_{t|t} = -BA^{-1}\pi_{t|t}.$$

The central bank acts to neutralize the effect of estimated demand shocks on $\pi_{t|t}$ and $y_{t|t}$ so that under the optimal Taylor rule, only the effects of cost-push shocks (i.e., "supply shocks") remain. As a consequence, the central bank's best estimates of current inflation and output are perfectly negatively correlated.[13]

2.5.2 Does the Federal Reserve Follow an Optimal Taylor Rule?

If the central bank followed an optimal Taylor rule (TR2) and the model were correct, then expected inflation and the expected output gap should be perfectly negatively correlated. To see whether this prediction holds in the data, we examine the Fed's real time forecasts of current quarter inflation and the output gap. These forecasts are produced by the Federal Reserve staff before every meeting of the Federal Open Market Committee and made publicly available with a five year lag by Federal Reserve Bank of Philadelphia. We take the Fed's contemporaneous estimate of Core CPI inflation as the empirical analogue of $\pi_{t|t}$.

Figure 2.5 plots the current-quarter forecasts of Core CPI inflation and the output gap from 1987Q3 - 2007Q4. In the figure, 'initial' refers to the forecast of the output gap produced for the first quarterly meeting and 'revised' refers to the forecast for the second meeting. Panel A shows the unfiltered series. There is no obvious relationship between the two series in the figure. The sample correlation between the initial output gap estimate and

[13]This result technically requires that the Taylor rule coefficients approach infinity at asymptotically the same rate so that their ratio converges to a strictly positive and finite value at the optimum.

the estimate of Core CPI inflation is -0.09 (the correlation changes to -0.1 if we use the revised estimate). In contrast, under the optimal Taylor rule, the correlation between the two series should be -1.00.

One concern with using the raw series to compute this correlation is that much of the variation in inflation is due to a steady declining trend since the mid 1980's. In Panel B, we show HP-filtered time series for both the estimated output gap and estimated inflation. (We use the standard quarterly smoothing parameter of 1600). In this figure, the estimates of inflation and the output gap have a modest positive correlation. The correlation coefficient is 0.35 for the initial output gap estimate and 0.29 for the revised estimate.

In neither case is the correlation close to the prediction of negative one. When interpreted through the lens of our model it appears that the Fed is underreacting to its own estimates of inflation and the gap.

2.5.3 Signal Extraction and Interest Rate Smoothing

There is a close connection between the Taylor rule given by (TR2) and interest rate smoothing. Interest rate smoothing can be captured by a policy rule of the form

$$i_t = \rho + \phi_\pi \pi_t^m + \phi_y y_t^m + \nu i_{t-1}. \tag{TR3}$$

In this specification the central bank sets the interest rate as a function of measured inflation, measured output, and the lagged interest rate.[14] The parameter ν governs the extent to which the central bank anchors its current policy with the interest rate from the previous quarter.

Taylor rules with interest rate smoothing are similar to Taylor rules based on filtered output and inflation. In fact, there are special cases in which the two rules exactly coincide.

[14] See, for example, *Sack and Wieland* (2000), and *Rudebusch* (2006) for discussions of interest rate smoothing.

To see this, consider the model in the previous section given by equations (2.1) to (2.6) and (TR2), in which the central bank observes $\pi_t^m = \pi_t + m_t^\pi$ and $y_t^m = y_t + m_t^y$ and uses the Kalman filter to estimate the output gap and inflation. Suppose further that there are no cost-push shocks ($V[u_t] = 0$) and that the persistence of the measurement error shocks is zero ($\varrho_{my} = \varrho_{m\pi} = 0$). For this special case, we have the following result.

Proposition 5: *Given the assumptions above, for any Taylor rule (TR2) with coefficients ψ_y and ψ_π, there exist coefficients $\{\tilde{\phi}_y, \tilde{\phi}_\pi, \tilde{\nu}\}$ such that the policy rule (TR3) generates the same equilibrium paths for all variables.*

Proposition 5 states that in the special case where there are no cost-push shocks and measurement error is i.i.d., the Taylor rule in which the central bank responds to the estimates $y_{t|t}$ and $\pi_{t|t}$ can be implemented exactly by a Taylor rule with interest rate smoothing of the form (TR3).[15]

While exact equivalence between interest rate smoothing and Taylor rules of the form (TR2) requires strong assumptions, the spirit of Proposition 5 extends to very general settings. As long as the central bank follows Taylor rule (TR2) and forms its estimates $y_{t|t}$ and $\pi_{t|t}$ using the Kalman filter, the interest rate will change gradually because the estimates of the output gap and inflation change gradually. In such a setting, an econometrician would continue to conclude that lagged interest rates play an important role in shaping policy. To demonstrate this claim quantitatively, we again consider the model consisting of equations (2.1) to (2.6), the central bank's informational constraints described above and Taylor rule (TR2). When we estimate the interest rate smoothing rule (TR3) on data simulated from this model using the benchmark calibration summarized in Table 2.1, we obtain a smoothing parameter ν of about 0.7.

[15] *Aoki* (2003) reaches a similar conclusion in a setting where the central bank optimizes under discretion.

2.6 Numerical Evaluation

Here we use a calibrated numerical model to quantitatively compare the policy rules (TR1), (TR2), and (TR3). All numerical results are based on the calibration discussed in Section 4 and summarized in Table 2.1. For each rule, we compute the optimal coefficients and evaluate the value of the objective function as well as the variance of inflation and output.

Table 2.2 reports the value of the objective function (equation (2.8)), as well as the associated variance of inflation and the output gap for all three policy rules evaluated at their optimal coefficients. For rules (TR1) and (TR3) the optimal coefficients are shown in the bottom rows of the table. Not surprisingly, the simple Taylor rule in which the central bank responds to measured inflation and measured output (equation (TR1)) is inferior to both alternatives. For our baseline calibration, Taylor rule (TR2) outperforms the interest rate smoothing rule (TR3) though we note that it is possible to construct cases in which optimal interest rate smoothing outperforms Taylor rule (TR2).[16]

Figures 2.6 and 2.6 report impulse response functions for the four structural shocks $(r_t^e, u_t, m_t^y, m_t^\pi)$. For each case, we consider a Taylor rule which is optimal given the specification (TR1), (TR2) or (TR3) as indicated. The top panels of Figure 2.6 show the impulse response functions of the output gap, inflation, and the interest rate to a shock to the efficient rate r_t^e (a "demand" shock). When the central bank follows the simple Taylor rule (TR1), the output gap is remarkably close to zero. However, inflation jumps up substantially and only slowly converges back to its steady state value of zero. The history dependent rules

[16]Consider a calibration in which all of the structural shocks have no persistence. In this case, the filtered estimates $\pi_{t|t} and y_{t|t}$ are functions of only current observationsπ_t^m and y_t^m. As a consequence, the optimized value of the objective is the same for the simple Taylor rule (TR1) and for the Taylor rule with filtered data (TR2). However, since a central bank following (TR3) can choose an additional parameter it must be able to weakly improve on the restricted rules. The interest rate smoothing parameter allows the central bank some ability to commit to future actions which is often a feature of the globally optimal policy. See, e.g., **?**, *Woodford* (1999) and *Rotemberg and Woodford* (1998).

(TR2) and (TR3) are superior in this regard.

Impulse response functions for the cost-push shock are shown in the bottom panels of Figure 2.6. On impact, the output gap is largest for the simple Taylor rule (TR1). No further striking differences of the paths of output and inflation are revealed for the different policy rules.

As in our baseline calibration, *Orphanides* (2003) estimated that measurement error in the output gap is highly persistent. The top panels of Figure 2.7 show that this persistence is reflected in the dynamic reaction to a noise shock. When the central bank solves a signal extraction problem and sets the interest rate according to Taylor rule (TR2), it learns gradually that the shock must have been measurement error. Although the output gap and inflation drop immediately after the shock, optimal state estimation reveals relatively quickly that the shock is likely noise and the central bank quickly guides the economy back to steady state. When the central bank follows rules (TR1) or (TR3) instead, the persistence of the measurement error shock causes the economy to go through a protracted period of low inflation.

Not surprisingly, signal extraction is particularly important when measurement error is persistent. Because our baseline calibration features no persistence for measurement error in inflation, when the economy experiences such a shock, there are only slight differences in output and inflation across the policies. This is shown in the bottom panels of Figure 2.7. While the impact response of the output gap is largest for the simple Taylor rule (TR1), the lack of history dependence implies a return to the steady state in a single period. Rules (TR2) and (TR3) show a somewhat slower convergence rate.

2.7 Conclusion

One of the Taylor rule's most appealing features is its simplicity: the central bank's behavior is characterized by two parameters only. In this paper we analyze optimal Taylor rules in standard New Keynesian models. In the absence of measurement error, activist monetary policy is costless and the optimal Taylor rule coefficients are often infinite. When inflation and output are measured with error, the optimal Taylor rule coefficients are finite. If the central bank instead sets the interest rate as a function of estimated output and inflation then the optimal coefficients on estimated inflation and the estimated gap are again infinite.

Optimal monetary policy in the model with signal extraction implies a strong negative correlation between estimated inflation and estimated output. In contrast, data on the Federal Reserve's estimates of current inflation and the output gap exhibit either zero correlation or a modest positive correlation depending on whether or not they are filtered. This suggests that the Fed is insufficiently aggressive in responding to deviations from its targets.

Signal extraction on the part of the central bank also introduces behavior which mimics history dependence in monetary policy. Because the central bank's beliefs about the state of the economy are updated gradually, the interest rate also changes gradually and observed policy actions will resemble interest rate smoothing even if the central bank is reacting only to current estimates of inflation and the output gap.

The results in this paper are not meant to be the final word on optimal Taylor rules in New Keynesian models but rather to provide a simple benchmark for optimal Taylor policy rules in more realistic and complex economic models. Such models would likely include wage rigidity in addition to price rigidity and would almost surely include durable consumer and investment goods. While we anticipate that optimal Taylor rules will differ in more articulated models, we also suspect that many of the properties we derived for the basic New Keynesian model will carry over to more realistic environments.

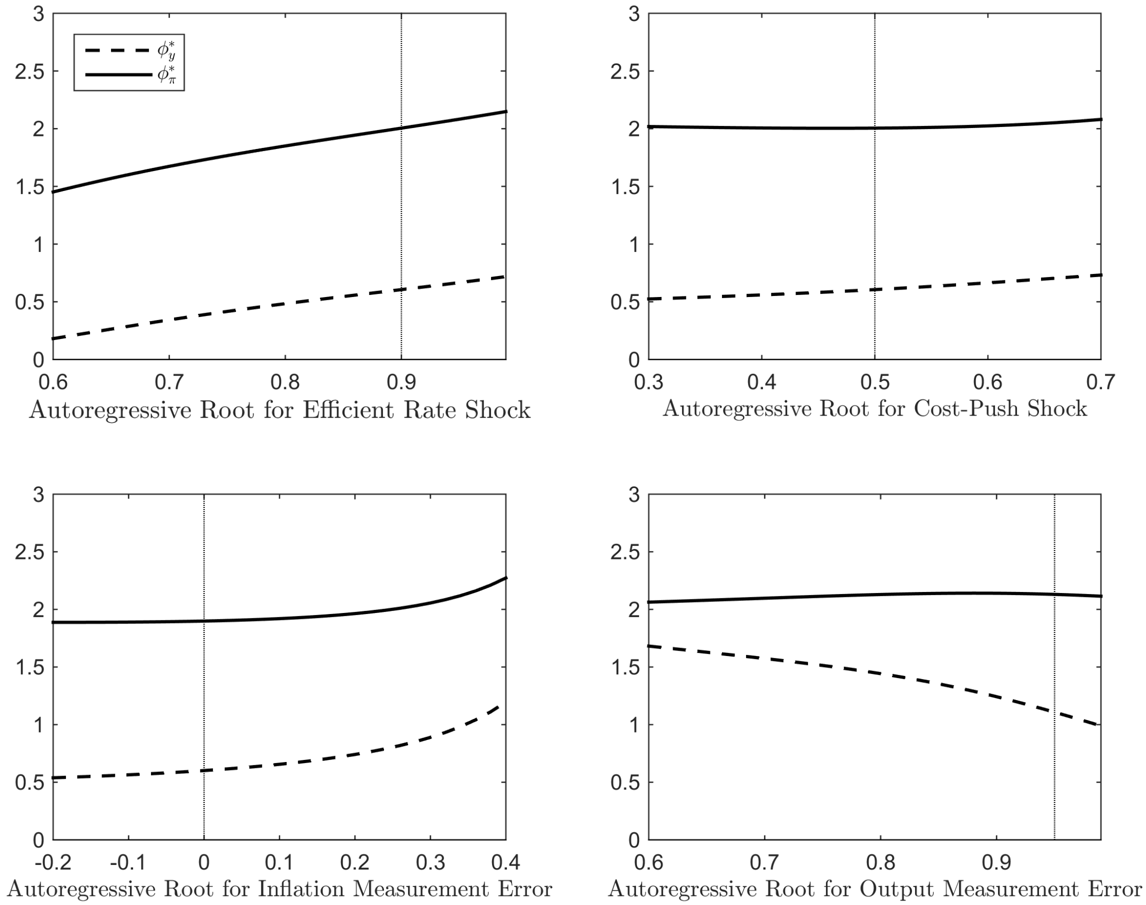
Table 2.1: Baseline Calibration

Parameter	Value	Description
β	0.99	Discount factor
σ	1	Coefficient of relative risk aversion
θ	2/3	Poisson rate of price stickiness
η	1	Labor supply elasticity
$(\varrho_r, \varrho_u, \varrho_{m\pi}, \varrho_{my})$	0.9 ,0.5 ,0 ,0.95	Persistence of shock processes
SD$[\varepsilon^r]$	0.22	Standard deviation (SD) of innovation to efficient rate
SD$[u]$	0.43	SD of innovation to the cost-push shock
SD$[\varepsilon^{m\pi}]$	0.50	SD of innovation to measurement error in inflation
SD$[\varepsilon^{my}]$	0.66	SD of innovation to measurement error in the gap
α	1	Weight on inflation in the policy objective function

Table 2.2: Evaluation of Policy Rules

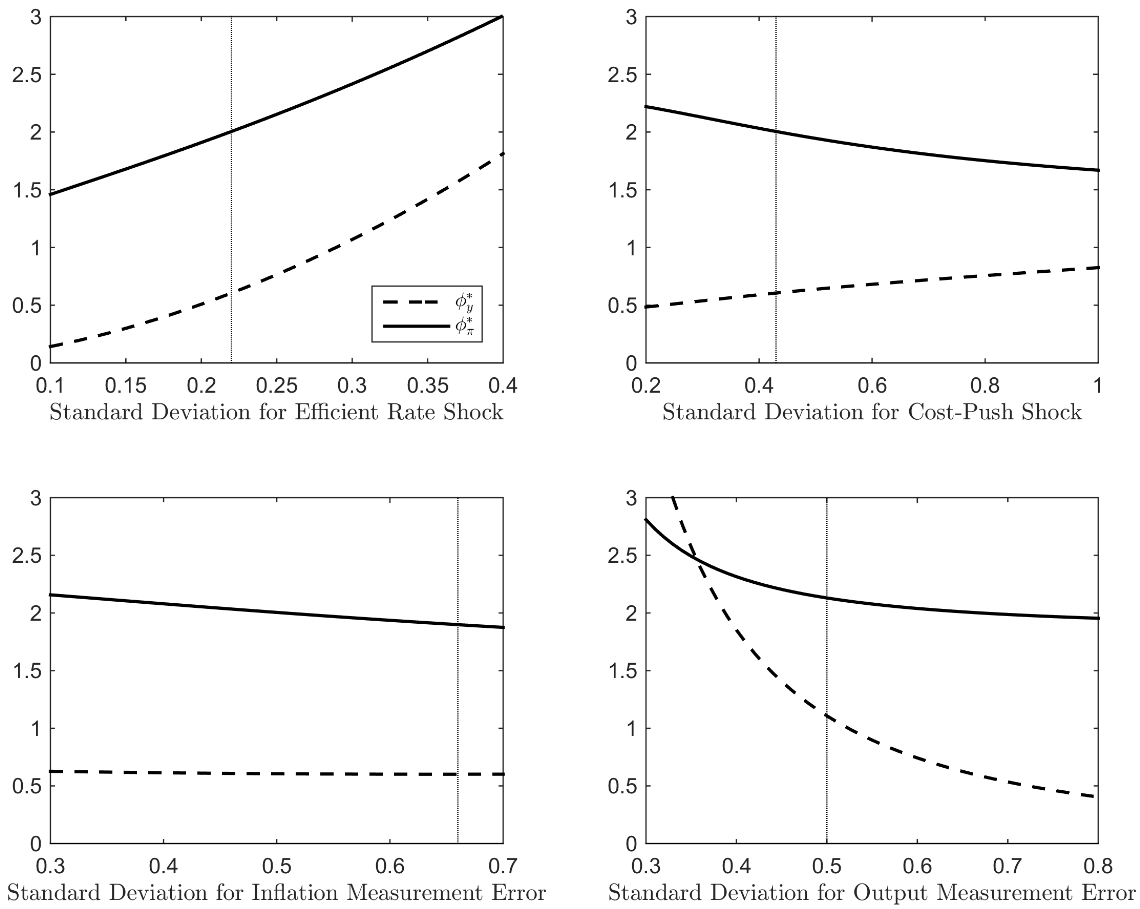
	Standard Taylor Rule (TR1)	Taylor Rule with signal extraction (TR2)	Interest Rate Smoothing(TR3)
Criterion (8)	1.56	1.03	1.16
V$[y]$	1.11	0.69	0.86
V$[\pi]$	0.45	0.34	0.30
ϕ_y^*	0.61	-	0.27
ϕ_π^*	2.00	-	0.42
ν^*	-	-	1.08

Figure 2.1: Parametric Variations in Shock Persistence



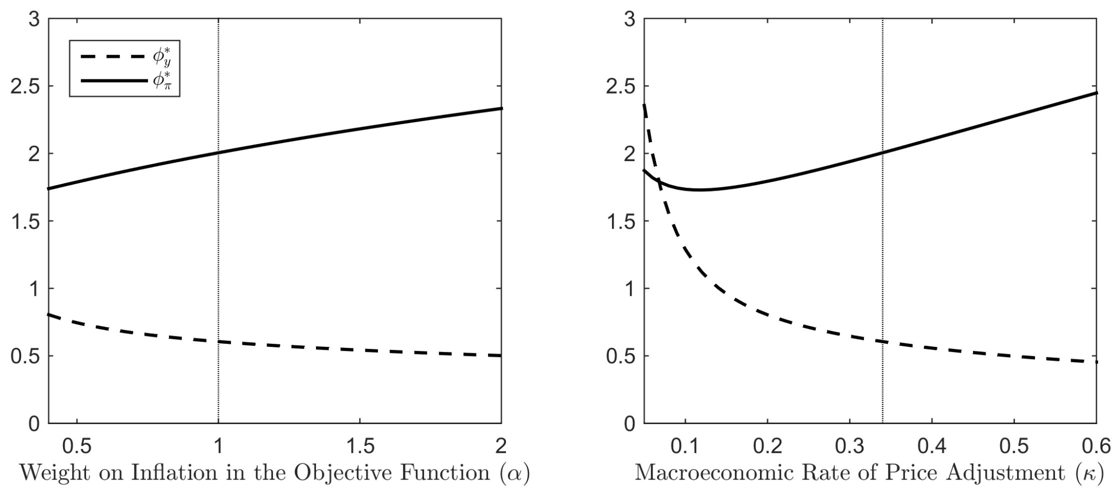
Notes: Each panel plots the optimal Taylor rule coefficients for inflation (heavy solid line) and the output gap (heavy dashed line). The vertical dashed line indicates the parameter value in the baseline calibration. For all cases, as the autoregressive coefficient changes, the innovation variance is adjusted so as to hold constant the variance of the overall process. All other parameters are held constant at the baseline level.

Figure 2.2: Parametric Variations in Shock Variance



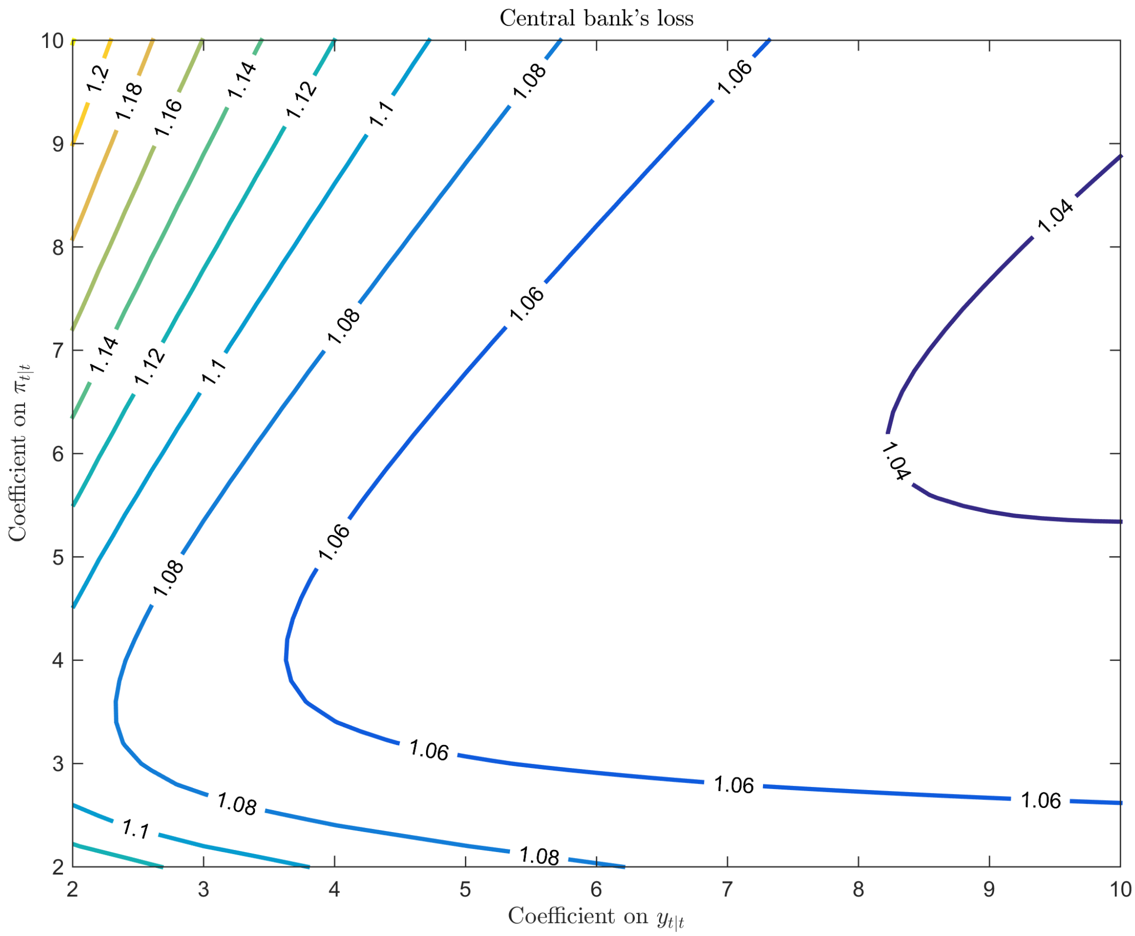
Notes: Each panel plots the optimal Taylor rule coefficients for inflation (heavy solid line) and the output gap (heavy dashed line). The vertical dashed line indicates the parameter value in the baseline calibration. All other parameters are held constant at the baseline level.

Figure 2.3: Parametric Variations in Inflation Weight and Rate of Price Adjustment



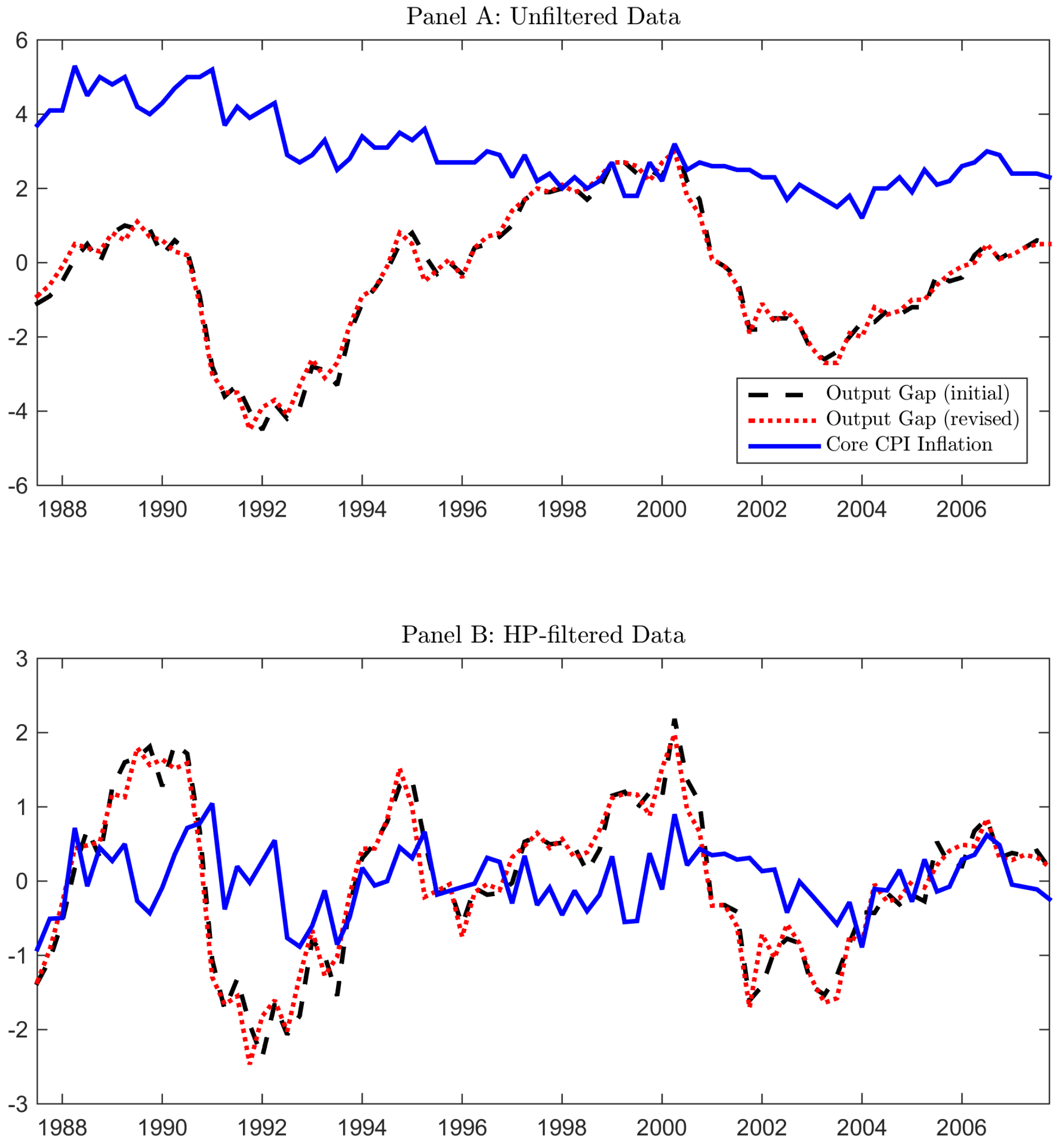
Notes: Each panel plots the optimal Taylor rule coefficients for inflation (heavy solid line) and the output gap (heavy dashed line). The vertical dashed line indicates the parameter value in the baseline calibration. All other parameters are held constant at the baseline level.

Figure 2.4: Level curves of the central bank's loss function



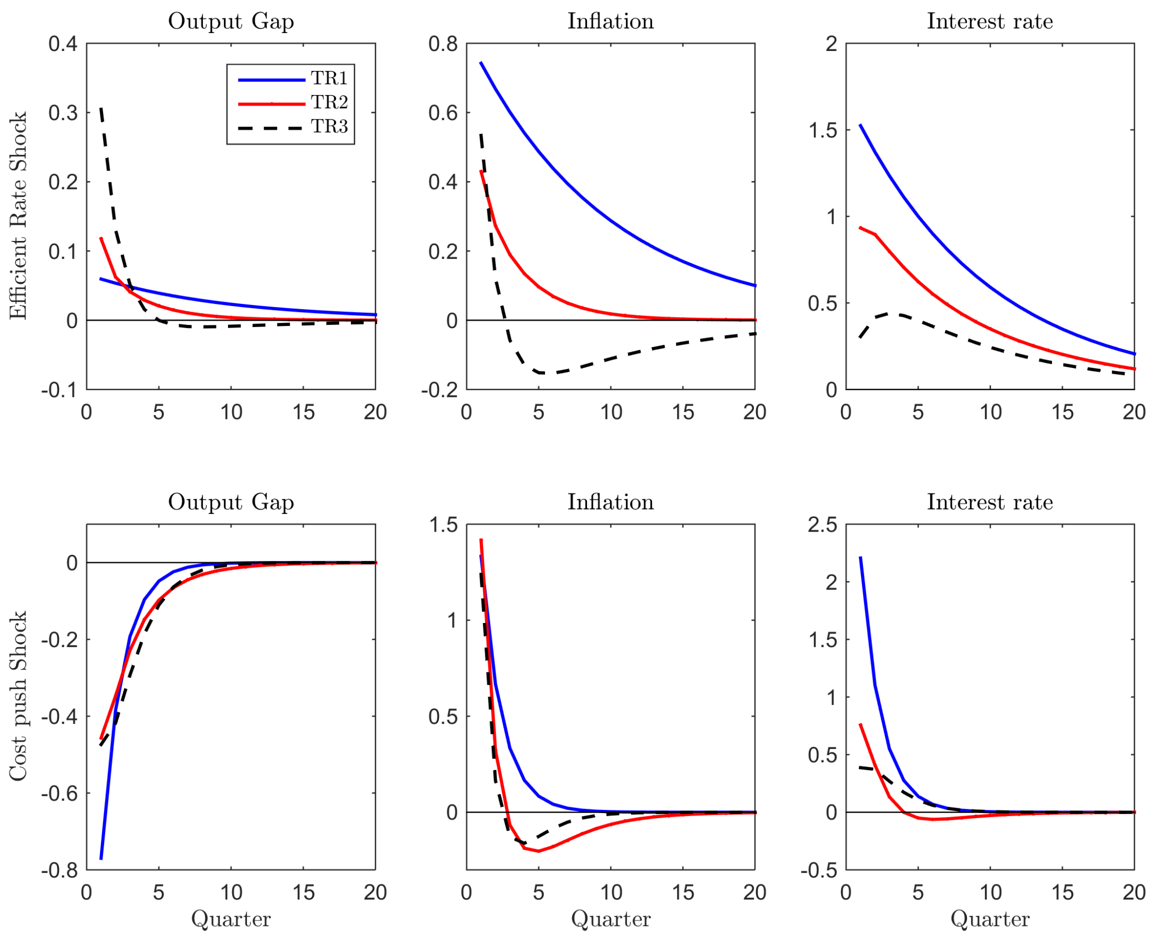
Notes: The figure uses the baseline calibration summarized in Table 2.1.

Figure 2.5: Inflation and output gap estimates of the Fed



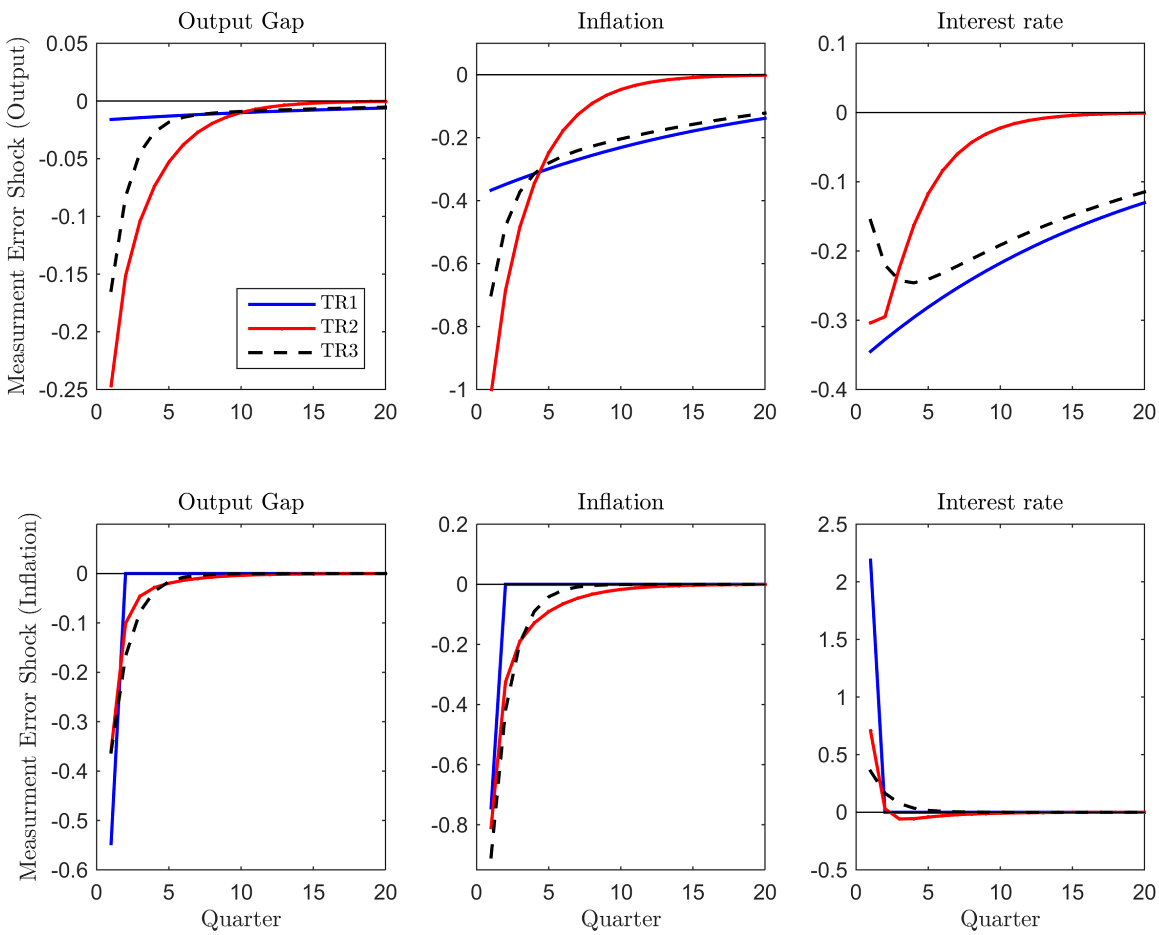
Notes: The graphs show the current-quarter estimates of Core CPI inflation and the output gap. All series in Panel B are HP-filtered with a smoothing parameter of 1600. The Federal Open Market Committee meets twice per quarter and two estimates for the output gap are available. Initial refers to the estimate for the first and revised to the estimate for the second meeting.
Source: Federal Reserve Bank of Philadelphia.

Figure 2.6: Impulse Response Functions



Notes: All shocks are one standard deviation in size. The unit of the output gap is percentage deviations from its steady state value of zero. The units of inflation and the interest rate are annualized percentage points.

Figure 2.7: Impulse Response Functions



Notes: All shocks are one standard deviation in size. The unit of the output gap is percentage deviations from its steady state value of zero. The units of inflation and the interest rate are annualized percentage points.

CHAPTER III

Input Linkages and the Transmission of Shocks: Firm-Level Evidence from the 2011 Tōhoku Earthquake

with Aaron Flaaen and Nitya Pandalai Nayar

3.1 Introduction

The spillover effects of trade and financial linkages has been a preeminant topic in international economics in recent decades. The large expansions in trade and foreign direct investment (FDI) in the past twenty years have generated much discussion on whether they increase volatility (*di Giovanni and Levchenko*, 2012), increase comovement (*Frankel and Rose*, 1998; *Burstein et al.*, 2008) or lead to less diversified production and specialization (*Imbs*, 2004). Identifying the micro-foundations underlying the role of these linkages in the increased interdependence of national economies is challenging. Advanced economies are highly connected, and most variables influenced by any candidate mechanism are often correlated with other developments in the source and destination countries. There is often little in the way of exogenous variation to isolate any particular mechanism from a host of confounding factors. Moreover, the requisite data to examine these issues at the necessary detail and disaggregation have been, until recently, unavailable.

This paper provides empirical evidence for the cross-country transmission of shocks via the rigid production linkages of multinational firms. The principal mechanism at work is not new; the idea of input-output linkages as a key channel through which shocks propagate through the economy dates back to at least *Leontief* (1936) or *Hirschmann* (1958). Two advances in this paper permit a new quantitative evaluation of the nature and magnitude of these linkages. First, we utilize a novel dataset that, for the first time, links restricted U.S. Census Bureau microdata to firms' international ownership structure. This information permits a forensic focus on particular firms and their underlying behavior. Second, we utilize the March 2011 Tōhoku earthquake and tsunami as a natural experiment of a large and exogenous shock disrupting the production linkages originating from Japan.

We study the role of imported intermediate inputs in the transmission of this shock to the United States economy. Because disruptions to imports of final goods would be unlikely to affect U.S. production, we develop a new methodology for isolating firm-level imports of intermediate inputs. We show that the U.S. affiliates of Japanese multinationals are the most natural source of this transmission, due to their high exposure to imported intermediates from Japan. The scope for shocks to these imported inputs to pass through and affect the firm's U.S. production depends on how substitutable they are with inputs from alternative sources. In other words, the role of imported inputs in the transmission of shocks is governed by the elasticity of substitution with respect to domestic factors of production.

We estimate this elasticity using the relative magnitudes of high frequency input and output shipments in the months following the Tōhoku earthquake/tsunami. This proceeds in two steps. First, reduced form estimates corresponding to Japanese multinational affiliates on average show that output falls, without a lag, by a comparable magnitude to the drop in imports. These results suggest a near-zero elasticity of imported inputs. Second, we structurally estimate a firm-level production function that allows for substitution across different types of inputs. The structural estimation procedure we use is uniquely tailored to the experiment. In an initial period prior to the Tōhoku disruption, we infer information on the firm's productivity and optimal input mix. Then, applying this production function to the period of the disruption, we estimate the elasticity parameters based on how changes in the firm's input mix translate into changes in output.

This estimation strategy has a number of attractive features. Most importantly, it relies on very few assumptions. Direct estimation of the production function circumvents the many difficulties associated with specifying a firm's optimization problem in the period after the shock. Second, it yields transparent parameter identification. This is an advantage over traditional estimation strategies as it does not suffer from omitted variables and endogeneity concerns arising from correlated shocks. Third, it allows for the estimation across different subgroups of firms.

The structural estimates are broadly in agreement with the results from our reduced form exercise. For Japanese multinationals, the elasticity of substitution across material inputs is 0.2 and the elasticity between material inputs and a capital/labor aggregate is 0.03. For non-Japanese firms using inputs from Japan, the estimates of the elasticity of substitution across material inputs are somewhat higher at 0.42 to 0.62. While the high cost share and particularly low elasticity for Japanese affiliates explains their predominant contribution to the direct transmission of this shock to the U.S., the elasticity estimates for non-Japanese firms are still substantially lower than typical estimates used in the literature. We argue that the substantial share of intra-firm intermediate trade implies greater complementarities in aggregate trade than is currently recognized.

There are a number of important implications for such low values of the elasticity of substitution. This parameter appears in various forms in a wide span of models involving the exchange of goods across countries. As discussed by *Backus et al.* (1994) and *Heathcote* *and Perri* (2002) among others, this parameter is critically important for the behavior of these models and their ability to match key patterns of the data. Prior estimates of this parameter were based on highly aggregated data that naturally suffered from concerns about endogeneity and issues of product composition.[1] Reflecting the uncertainty of available estimates for the elasticity of substitution, it is a common practice to evaluate the behavior of these models along a wide range of parameter values.

It is well known that a low value for this parameter (interpreted as either substitution between imported and domestic goods in final consumption or as intermediates in production) improves the fit of standard IRBC models along several important dimensions. In particular, the elasticity of substitution plays a role in two highly robust failings of these models: i) a terms of trade that is not nearly as variable as the data, and ii) a consumption comovement that is significantly higher than that of output, whereas the data show the opposite relative ranking.[2]

To understand the relationship between the elasticity and comovement, it is helpful to recall that these models generate output comovement by inducing synchronization in factor supplies, a mechanism that by itself generally fails to produce the degree of comovement seen in the data. Complementarities among inputs together with heterogeneous input shocks will generate direct comovement in production, augmenting the output synchronization based on factor movements. *Burstein et al.* (2008) show that a low production elasticity of substitution between imported and domestic inputs reduces substitution following relative price movements, and thereby increases business cycle synchronization.[3] It is also relatively straightforward to see how a lower elasticity increases volatility in the terms of trade. When

[1] For a very useful compendium of this research from this era, see *Stern et al.* (1976). More recently, work by *Halpern et al.* (2011) and *Goldberg et al.* (2010) demonstrate that materials inputs from foreign countries are imperfectly substitutable with domestic inputs for Hungary and India respectively.

[2] Due to the robust nature of these shortcomings, *Backus et al.* (1995) refer to them as the "price anomaly" and "quantity anomaly" respectively.

[3] Although they do not estimate this parameter, the value they advocate (0.05) is indeed close to our estimates.

two inputs are highly complementary, deviations from the steady state mix are associated with large changes in their relative prices. In the words of *Heathcote and Perri* (2002, page 621): "greater complementarity is associated with a larger return to relative scarcity."

The estimates in this paper have implications for the role of trade in firm-level and aggregate volatility. Other research has argued that firms can diversify risk arising from country specific shocks by importing (*Caselli et al.* (2014)) or that firms with complex production processes of several inputs are less volatile as each input matters less for production (*Koren and Tenreyro* (2013)). On the other hand, there is a well-established fact that complementarities and multi-stage processing can lead to the amplification of shocks as in *Jones* (2011) and *Kremer* (1993). We discuss the potential for measured amplification in our context in Section 3.5.

This paper is also a contribution to the empirical evidence on the role of individual firms in aggregate fluctuations, emanating from the work of *Gabaix* (2011). Other related evidence comes from *di Giovanni et al.* (2014), who use French micro-data to demonstrate that firm-level shocks contribute as much to aggregate volatility as sectoral and macroeconomic shocks combined. The so-called granularity of the economy is very much evident in our exercise; though the number of Japanese multinationals is small, they comprise a very large share of total imports from Japan, and are arguably responsible for a measurable drop in U.S. industrial production following the Tōhoku earthquake (see Figure 3.3).

The strong complementarity across material inputs implies that non-Japanese input use falls nearly proportionately, thereby propagating the shock to other upstream (and downstream) firms in both the U.S. economy and abroad. Many suppliers were thus indirectly exposed to the shock via linkages with Japanese affiliates that had i) high exposure to Japanese inputs and ii) a rigid production function with respect to other inputs. Network effects such as these can dramatically magnify the overall transmission of the shock (both across countries and within). And while such effects are commonly understood to exist, this paper provides unique empirical evidence of the central mechanisms at work.

As is the case with most research based on an event-study, some care should be taken in generalizing the results to other settings. Although we have already highlighted the aggregate implications of the effects we estimate, one might worry that the composition of Japanese trade or firms engaged in such trade is not representative of trade linkages more broadly. We believe the results we obtain are informative beyond the context of this particular episode for two reasons. First, the features of Japanese multinationals that are underlying the transmission of this shock are common to all foreign multinational affiliates in the United States.[4] Second, estimates corresponding to all firms in our sample also exhibit substantial complementarities, and as a whole these firms account for over 70 percent of total U.S. manufacturing imports.

The next section describes the empirical strategy and data sources used in this paper, section 3.3 presents reduced form evidence in support of a low production elasticity of imported inputs for Japanese multinational affiliates. In Section 3.4, we expand the scope of parameters we identify with a structural model of cross-country production linkages. We estimate the parameters of this model across several different subgroups. Section 3.5 discusses the implications of these estimates, and details a number of checks and robustness exercises. The final section offers concluding thoughts.

3.2 Empirical Strategy and Specification

This section outlines the empirical approach of using an event-study framework surrounding the 2011 Tōhoku event to estimate the production elasticity of imported inputs. We discuss the relevant details of this shock, document the aggregate effects, and then outline the empirical specification for the firm-level analysis.

[4] Intra-firm trade accounts for a large majority of the trade of Japanese affiliates. More generally, the intra-firm share of imported intermediates for all foreign affiliates in the U.S. is 71 percent.

3.2.1 Background

The Tōhoku earthquake and tsunami took place off the coast of Northeast Japan on March 11, 2011. It had a devastating impact on Japan, with estimates of almost twenty thousand dead or missing (*Schnell and Weinstein* (2012)) and substantial destruction of physical capital. The magnitude of the earthquake was recorded at 9.0 on the moment magnitude scale (M_w), making it the fourth largest earthquake event recorded in the modern era.[5] Most of the damage and casualties were a result of the subsequent tsunami that inundated entire towns and coastal fishing villages. The effects of the tsunami were especially devastating in the Iwate, Miyagi, and Fukushima prefectures. The Japanese Meteorological Agency published estimates of wave heights as high as 7-9m (23-29ft), while the Port and Airport Research Institute (PARI) cite estimates of the maximum landfall height of between 7.9m and 13.3m (26-44ft).

Figure 3.1 shows the considerable impact of the Tōhoku event on the Japanese economy. Japanese manufacturing production fell by roughly 15 percentage points between February and March 2011, and did not return to trend levels until July. Much of the decline in economic activity resulted from significant power outages that persisted for months following damage to several power plants – most notably the Fukushima nuclear reactor.[6] Further, at least six Japanese ports (among them the Hachinohe, Sendai, Ishinomaki and Onahama) sustained significant damage and were out of operation for more than a month, delaying shipments to both foreign and domestic locations. It should be noted, however, that the largest Japanese ports (Yokohama, Tokyo, Kobe) which account for the considerable majority of Japanese

[5] Since 1900, the three earthquakes of greater recorded magnitude are: the 1960 Great Chilean earthquake (magnitude 9.5), the 1964 Good Friday earthquake in Prince William Sound, Alaska (magnitude 9.2); and the 2004 Sumatra-Andaman earthquake (magnitude 9.2).

[6] For precautionary reasons, all nuclear power plants were immediately shut down following the earthquake, and remained largely offline until 2014 or later. Because the electricity infrastructure exists on two separate grids (a 60Hz to the south and west, and 50Hz to the north and east), the reduction in power supply in Northeast Japan was not easily remedied, and power outages persisted for months.

trade, re-opened only days after the event.

As expected, the economic impact of the event was reflected in international trade statistics, including exports to the United States. Figure 3.2 plots U.S. imports from Japan around the period of the Tōhoku event, with imports from the rest of the world for comparison. The large fall in imports occurs during the month of April 2011, reflecting the several weeks of transit time for container vessels to cross the Pacific Ocean. The magnitude of this drop in imports is roughly similar to that of Japanese manufacturing production: a 20 percentage point drop from March to April, with a full recovery by July 2011.

More striking is the response of U.S. industrial production in the months following the event. Figure 3.3 demonstrates that there is indeed a drop in U.S. manufacturing production in the months following the Japanese earthquake. Although the magnitudes are obviously much smaller — roughly a one percentage point drop in total manufacturing and almost two percentage points in durable goods — the existence of a measurable effect is clear.[7]

Though tragic, the Tōhoku event provides a glimpse into the cross-country spillovers following an exogenous supply shock. This natural experiment features many characteristics that are advantageous for this type of study. It was large and hence measurable, unexpected, and directly affected only one country. The shock was also short-lived, which rules out immediate supplier restructuring and allows for an estimate of the elasticity for a given supply chain.[8] On the other hand, the short duration of the shock presents a challenge for measurement as it limits the available datasets with information at the required frequency. We utilize a novel firm-level dataset to uncover the mechanisms at work behind the transmission of this shock.

[7]At the level of total U.S. GDP, both Deutsche Bank and Goldman Sachs revised 2nd quarter U.S. estimates down by 50 basis points explicitly due to the events in Japan.

[8]It also rules out large balance sheet effects that would make differential credit conditions an operative feature.

3.2.2 Data

Several restricted-use Census Bureau datasets form the core of our firm-level analysis. The Longitudinal Business Database (LBD) collects the employment, payroll, and major industry of all establishments operating in the United States, and is maintained and updated as described by *Jarmin and Miranda* (2002). Longitudinal linkages allow the researcher to follow the establishment over time, and the annual Company Organization Survey (COS) provides a mapping from establishments to firms. All of the analysis in this paper will be at the firm-level.

The Longitudinal Foreign Trade Transactions Database (LFTTD), which links individual trade transactions to firms operating in the United States. Assembled by a collaboration between the U.S. Census Bureau and the U.S. Customs Bureau, the LFTTD contains information on the destination (or source) country, quantity and value shipped, the transport mode, and other details from point-of-trade administrative documents. Importantly for this study, the LFTTD includes import and export trade transactions at a *daily* frequency, which is easily aggregated to monthly-level trade flows. A number of important papers have utilized this resource, such as *Bernard et al.* (2007) and *Bernard et al.* (2006).

We utilize two novel extensions to this set of Census data products. First, a new link between a set of international corporate directories and the Business Register (BR) of the Census Bureau provides information on the international affiliates of firms operating in the United States. These directories provide information, for the first time, to identify those U.S. affiliates part of a foreign parent company, as well as those U.S. firms with affiliate operations abroad. This information is an important resource for identifying the characteristics of U.S. firms affected by the Tōhoku event. For information on these directories and the linking procedure used, please see Appendix C.2.1.

The second novel data resource is a system to classify firm-level import transactions as intermediate or final goods. Although intermediate input trade represents as much as two-thirds of total trade (see *Johnson and Noguera* (2012)), the LFTTD does not classify a trade transaction based on its intended use. To overcome this limitation, we use information on the products produced by U.S. establishments in a given industry to identify the set of products intended for final sale *for that industry.*[9] The remaining products are presumably used by establishments in that industry either as intermediate inputs or as capital investment. Details on this classification procedure are available in Appendix C.2.2. In the aggregate, this firm-level classification procedure yields estimates of the intermediate share of trade that are consistent with prior estimates: 64 percent of manufacturing imports are classified as "intermediates" in 2007.

Finally, we utilize geographic information on the severity of the earthquake/tsunami that is compiled by the U.S. Geological Survey (USGS). By geocoding the Japanese addresses of firms with U.S. operations, we construct an earthquake intensity measure for each Japanese affiliate location. We then apply such information to the U.S. operations as a way to further measure the sample of firms plausibly affected by the shock. Please see Appendix C.2.3.2 for details. Figure 3.4 shows the geographic distribution of one such USGS measure — the modified mercalli index (MMI) — along with the geocoded affiliate locations.

The ideal dataset to evaluate the transmission of the Tōhoku event on U.S. firms would consist of high frequency information on production, material inputs, and trade, separated out by geographic and ownership criteria. Unfortunately, Census data on production and material inputs at the firm-level is somewhat limited. The Annual Survey of Manufacturers (ASM) contains such information, but at an annual frequency and only for a subset of manufacturing firms. On the other hand, firm-level trade information is available at a nearly daily frequency, and covers the universe of firms engaged in exporting/importing. For the

[9]Note that products intended for final sale for a given industry may still be used as intermediates for other firms in a different industry. Alternatively, such "final goods" can be sold directly to consumers for ultimate consumption.

purposes of characterizing the shock to firm-level imports of intermediate goods, the LFTTD (and supplements identified above) is ideal. There remain significant gaps in information on a firm's domestic input usage, a limitation we discuss in subsequent sections.

Because of the challenges of high-frequency information on firms' U.S. production, we utilize a proxy based on the LFTTD — namely the firm's exports of goods to North America (Canada and Mexico). The underlying assumption of this proxy is that all firms export a fixed fraction of their U.S. output to neighboring countries in each period. The advantage of this approach is the ability to capture the flow of goods at a specific point in time. There are few barriers to North American trade, and transport time is relatively short. Moreover, exporting is a common feature of these firms, of which exports to North America is by far the largest component. The obvious disadvantage of this approach is that it conditions on a positive trading relationship between firms in the U.S. and Canada/Mexico. We will assess the quality of this measure as a proxy for output in section 3.5.3.1.[10]

3.2.3 Basic Theory

Before moving to our firm-level analysis, it is useful to describe the basic theoretical structure of the features of firm-level production that we estimate. The transmission of shocks within a firm's production chain is governed by the flexibility of production with respect to input sourcing. Rather than model these complex networks directly, the literature typically summarizes this feature with the well-known elasticity of substitution within a C.E.S. production function. Our identification of this elasticity will rely on the relative impacts on output and imported inputs following the shock. To be concrete, consider the

[10]Another consideration with the use of this proxy is whether it more accurately reflects production or sales, as the two are distinct in the presence of output inventories. In our case, this depends on whether the inventories are held in the U.S. or Canada/Mexico. Without further evidence, we interpret the proxy to be capturing some mix between production and sales. The structural estimation in section 3.4 will allow for such a mix.

C.E.S. production function

$$x = \left[(1-\mu)^{\frac{1}{\psi}} \left[F_D \right]^{\frac{\psi-1}{\psi}} + \mu^{\frac{1}{\psi}} \left[IM \right]^{\frac{\psi-1}{\psi}} \right]^{\frac{\psi}{\psi-1}} \tag{3.1}$$

where output consists of combining a domestic bundle of factors F_D (e.g. capital and labor) with a foreign imported input IM. The parameter μ reflects the relative weight on the input IM in production, conditional on prices and a given elasticity value. Suppose the firm purchases its inputs in competitive markets with prices p_D and p_M, respectively, and sells its good at price p_x. Our approach in this section will be to estimate the parameter ψ governing the degree of substitution between these inputs, using information on the output elasticity with respect to imported inputs, $\frac{\partial \ln p_x x}{\partial \ln p_M M}$, in the months following the shock.

The first order conditions imply that

$$\frac{F_D^*}{IM^*} = \frac{1-\mu}{\mu} \left(\frac{p_M}{p_D} \right)^{\psi}, \tag{3.2}$$

where F_D^* and IM^* denote the optimal quantities of inputs. We would like to show the theoretical foundations underlying the intuitive result that a one-for-one drop in output with the fall in imported inputs implies an elasticity of zero. To do this, we make the following assumptions, all of which we will relax to some degree in the estimation framework in Section 3.4:

1. Imported inputs shipments are disrupted, such that the firm receives a suboptimally low quantity of IM: $IM < IM^*$;

2. The firm is unable to adjust domestic inputs F_D^* or its price p_x after learning that it receives IM;

3. The firm does not shut down.

Given these assumptions, the following result holds:

Result 1. *Under assumptions 1) to 3):*

$$\frac{\partial \ln p_x x}{\partial \ln p_M IM} = \frac{1}{1 + \left(\frac{IM^*}{IM}\right)^{\frac{\psi-1}{\psi}} \left(\frac{1-\mu}{\mu}\right) \left(\frac{p_M}{p_D}\right)^{\psi-1}} \in (0,1) \tag{3.3}$$

for any $\psi \in (0,\infty)$.

Proof. See Appendix C.1.1 for details. □

An immediate implication of this result is that the output elasticity is unity only when ψ approaches zero.[11] In this case $\left(\frac{IM^*}{IM}\right)^{\frac{\psi-1}{\psi}} \to 0$ (recall that $IM < IM^*$) and hence $\lim_{\psi\to 0} \frac{\partial \ln p_x x}{\partial \ln p_M IM} = 1$. Hence, observing a one-for-one drop in the value of output with the value of imported intermediates, we infer that ψ is close to zero. It is also straightforward to show that conditional on a value for $\psi \in (0,\infty)$, the output elasticity in (3.3) is increasing in the parameter μ. That is, conditional on a given drop in the imported input, a larger weight on this input leads to a larger percent response in output.

Our use of the natural experiment is critical for observing the effects of suboptimal input combinations (F_D^*, IM). To see this, suppose the firm could freely adjust F_D after learning it will receive $IM < IM^*$. Then, it would choose F_D such that $\frac{F_D}{IM} = \frac{F_D^*}{IM^*}$ and the firm would contract one-for-one with the drop in imports. It is a well-known fact that constant returns to scale production functions in competitive environments lead to indeterminate firm size. This has the implication that:

$$\frac{\partial \ln (p_x x)}{\partial \ln (p_M IM)} = \frac{\partial \ln (p_x x)}{\partial \ln (p_D F_D)} = \frac{\partial \ln (p_D F_D)}{\partial \ln (p_M IM)} = 1. \tag{3.4}$$

[11]There is a second case which we do not examine, where $\psi \to \infty$ and $p_M < p_D$ and thus the firm only uses IM. We discard this scenario because such a firm would not show up in our data (i.e. this case implies zero U.S. employment).

In this case it is not possible to learn anything about ψ from the joint behavior of output and the value of intermediate inputs. We provide evidence below that firms did not significantly adjust their domestic labor force following the disruption, so that a constant F_D is indeed a reasonable assumption in this simple framework. To be sure, there are a number of alternative frameworks where such behavior would not hold. We discuss some of these in Appendix A, and show that the mapping $\lim_{\psi \to 0} \frac{\partial \ln p_x x}{\partial \ln p_M IM} = 1$ is more general.

3.3 Reduced Form Evidence

This section will provide intuitive reduced-form evidence on the elasticity of substitution corresponding to the U.S. affiliates of Japanese multinationals. We discuss our strategy for understanding this elasticity via firm-behavior in the months following the Tōhoku event, and then report the results.

3.3.1 Framework

Our analysis of the production function (3.1) above demonstrates that a natural measure to evaluate the potential conduits of the Tōhoku shock to the United States would be the degree of reliance on Japanese imported inputs. This is best expressed as the cost share of inputs from Japan, and can be constructed in a Census year by taking a firm's Japanese imported inputs and dividing by all other inputs (which includes production worker wages and salaries, the cost of materials, and the cost of new machinery expenditures). Exposure to Japanese imported inputs is heavily concentrated among Japanese affiliates. In the year 2007, which is the closest available Census year, this cost share was nearly 22% on average for Japanese affiliates (see Table 3.1), compared to just 1% for other firms. For more detail on the heterogeneity across and within these firm groups, we construct a density estimate of such an exposure measure for the Japanese affiliates and non-Japanese multinationals. The results, shown in Figure 3.5, show little overlap between these distributions: there are few Japanese affiliates with low exposure to Japanese inputs, and few non-Japanese firms have substantial exposure.[12]

We now estimate the relative impacts on imported inputs and output for the Japanese affiliates as a group. To do this, we implement a dynamic treatment effects specification in which a firm is defined as being treated if it is owned by a Japanese parent company.[13] The effect on these firms can be inferred from the differential impact of the variable of interest relative to a control group, which soaks up common seasonal patterns and other demand-driven factors in the U.S. market. While there are a number of competing methodologies for this type of estimation, we use normalized propensity score re-weighting due to the relatively favorable finite-sample properties as discussed in *Busso et al.* (2014), as well as for its transparent intuition. Consistent estimation of the average treatment effect on the treated requires the assumption of conditional independence: the treatment/control allocation is independent of potential outcomes conditional on a set of variables. As the average Japanese firm differs considerably from other firms in the data, we use other multinational firms – both US and non-Japanese foreign- as our baseline control group prior to reweighting. To compute the propensity scores for reweighting, we control for size and industry, which ensures the control group has a similar industrial composition and size distribution as our treated sample.[14] Table 3.2 reports summary values for the sample, including statistics on the balancing procedure using the normalized propensity score.

[12]The exposure measure used in Figure 3.5 is from 2010 and does not include the cost of domestic material usage.

[13]We could have also used a threshold of Japanese input usage for the classification of treatment status. Doing so yields estimates that are very similar, which is due to the patterns evident in Figure 3.5. We have also tried conditioning on our geographic information (i.e. the firm-level Japanese MMI index) in defining a Japanese firm as being treated. The results are largely unchanged from those we report here, and for the sake of clarity we report results pertaining to the full sample.

[14]Using the predicted values (p) from the first stage regression, the inverse probability weights are $\frac{1}{1-p}$ for the control group and $\frac{1}{p}$ for the treated group. To normalize the weights such that the treated firms have weights equal to one, we then multiply each set of weights by p.

The magnitude of the shock for a representative Japanese multinational is captured by the effect on total imported intermediate products at a monthly frequency.[15] Including non-Japanese imported intermediates is important for applying the control group as a counterfactual, and the shares by source-country gives the necessary variation for identification: as shown in Table 3.2 the share of imported inputs from Japan is 70% of the total for Japanese firms and only 3.5% for non-Japanese multinationals. Let $V_{i,t}^M$ be the value of intermediate imports of firm i in month t, after removing a firm-specific linear trend through March 2011. We fit the following regression:

$$V_{i,t}^M = \alpha_i + \sum_{p=-4}^{9} \gamma_p E_p + \sum_{p=-4}^{9} \beta_p E_p \text{JPN}_{i,p} + u_{i,t} \tag{3.5}$$

where α_i are firm fixed-effects, γ_p are monthly fixed effects (with the indicator variables E_p corresponding to the calendar-months surrounding the event), and $u_{i,t}$ is an error term. The baseline sample will consist of January 2009 to December 2011. We denote March 2011 as t=0.

The β_p coefficients are of primary interest. The $\text{JPN}_{i,t}$ is an indicator variable equal to one if the firm is owned by a Japanese parent company. Interacting these indicator variables with each month of the panel allows for a time-varying effect of Japanese ownership on a firm's overall intermediate input imports, particularly during and after the Tōhoku event. The β_p coefficients will estimate the differential effect of the Tōhoku event on Japanese multinational affiliates in the U.S., compared to the control group of non-Japanese firms. A useful interpretation of the $\{E_p \text{JPN}_{i,p}\}$ variables is as a set of instruments that captures the exogeneity of imports during these months, reflecting the source-country share of imports from Japan as evident in Table 3.2. To evaluate the differential impact on production for Japanese firms, we simply replace the dependent variable in equation (3.5) with the firm's

[15] We consider Japanese and non-Japanese intermediate imports separately in section 3.4.

North American exports, denoted $V_{i,t}^{NA}$.

It is important to highlight that equation (3.5) is in levels. There are several reasons for doing so, as opposed to using log differences or growth rates. First, allowing for the presence of zeros is important when the data are at a monthly frequency, particularly given the magnitude of the shock to imports for Japanese firms. The second reason is more conceptual. Because we are interested in calculating the average effect of these firms that represents (and can scale up to) the aggregate impact on the U.S. economy, it is appropriate to weight the firms based on their relative size. The levels specification does exactly this: the absolute deviations from trend will be greater for the bigger firms and hence will contribute disproportionately to the coefficient estimates.[16] In section 3.4, we evaluate this framework with the results one would obtain when estimating the effect on a firm-by-firm basis.

In addition to the Conditional Independence Assumption highlighted earlier, the β_p coefficients are valid estimates of the mean effect for Japanese affiliates only in so far as the control group is not itself impacted by the shock. This Stable Unit Treatment Value Assumption (SUTVA) implies that general equilibrium effects or peer effects (e.g. strategic interaction) do not meaningfully effect the estimates. The share of imported inputs from Japan is low for the control group, and thus the shock is unlikely to have a measurable effect on imported inputs as a whole. We discuss strategic interaction in section 3.5.3.4.

3.3.2 Results: Total Manufacturing Sector

The top panel of Figure 3.6 plots the β_p coefficients from equation (3.5) for the months surrounding the Tōhoku event. Relative to the control group, there is a large drop in total intermediate input imports by Japanese firms in the months following the earthquake. The drop in intermediate inputs bottoms out at 4 million USD in $t = 3$ (June 2011) and the

[16] See Appendix C.3.1 for more discussion, as well as results obtained using other specifications. Importantly, in a reduced sample abstracting from zeros, a weighted regression using percentage changes directly yields estimates that are very close to those presented here.

point estimates do not return back to the pre-shock trend until month $t = 7$ (October 2011).

More interesting are the results from panel B of Figure 3.6, which looks for evidence of the production/sales impact of this shock on Japanese firms via their North American exports. The differential time-path of N.A. exports also exhibits a substantial drop following the Tōhoku event, hitting a trough of 2 million USD below baseline in $t = 2$ (May 2011). The standard errors, which are clustered at the firm level, are themselves interesting. As made clear via the 95-percent confidence bands on the point estimates of Figure 3.6, the standard errors increase dramatically in the months following the shock, a feature we interpret to reflect heterogeneous incidence and timing of the shocks (as well as the recoveries) for the Japanese multinationals.

To gain a sense of the average percentage drops of these two data series for Japanese multinationals as a group, we take the two plots of the differential dollar amounts from Figure 3.6 and divide by the average pre-shock level for these firms (see Table 3.2). The results, plotted jointly in Figure 3.7, show the fraction below pre-shock trend levels for these firms, on average. There is a remarkable correlation between these two series – whereby there is essentially a one-for-one drop in output for a given drop in intermediate imports. Using the mapping from Result 1, these reduced form results suggest a production function that is essentially Leontief in the imported input.

One potential concern with the interpretation of these results is separating out the intermediate input channel with other channels, such as a direct "productivity shock" affecting the U.S. operations of Japanese affiliates. Separating an ownership channel from an imported input channel is difficult due to lack of substantial overlap we identified above: few Japanese firms have low input exposure and few non-Japanese firms have high input exposure. In appendix C.3.2 we present results using a binary response model to disentangle the defining features of the import and output disruptions during this time.

3.4 Structural Estimation of Cross Country Input Linkages

The relative movements of imported inputs and output of Japanese multinational firms point to little substitutability of Japanese intermediate inputs. In this section we expand our analysis by structurally estimating the production function of firms affected by the Tōhoku shock. Unlike in the previous section, which used a set of instruments related to the differential import share of intermediates coming from Japan, this estimation relies on leveraging the high degree of exogenous variation in Japanese inputs coming from the Tōhoku event, while also fully specifying the production function under study. This estimation serves multiple purposes. First, it is reassuring to find elasticities that are consistent with the heuristic evidence implied by our reduced-form results, when imposing a conventional production function framework. Second, by imposing additional structure, we are able to distinguish two elasticities: one between Japanese material inputs and other material inputs, and another between an aggregate bundle of material inputs and domestic capital and labor. Finally, by using an estimation procedure not relying on a control group we can obtain separate estimates for Japanese and non-Japanese firms. The results corroborate the claim that the supply chains of Japanese and non-Japanese exhibit different degrees of rigidity.

The estimation procedure will utilize information from two distinct periods: the six months preceding and the six months following the March 11 event. The pre-period, which we denote by $\tau - 1$, yields information on the production function of the firm under profit-maximizing conditions. In the post-period, denoted τ, we do not impose that the firm is optimizing over its input use, due to the fact that shipments from Japan are to some extent beyond the control of the firm.

3.4.1 Framework

We assume that the firm's technology in any period t is given by the nested CES aggregate

$$x_{i,t} = \phi_i \left[\mu_i^{\frac{1}{\zeta}} \left(K_{i,t}^{\alpha} L_{i,t}^{1-\alpha} \right)^{\frac{\zeta-1}{\zeta}} + (1-\mu_i)^{\frac{1}{\zeta}} M_{i,t}^{\frac{\zeta-1}{\zeta}} \right]^{\frac{\zeta}{\zeta-1}}, \tag{3.6}$$

where

$$M_{i,t} = \left(\nu_i^{\frac{1}{\omega}} \left(m_{i,t}^{-J} \right)^{\frac{\omega-1}{\omega}} + (1-\nu_i)^{\frac{1}{\omega}} \left(m_{i,t}^{J} \right)^{\frac{\omega-1}{\omega}} \right)^{\frac{\omega}{\omega-1}}. \tag{3.7}$$

In this production function $x_{i,t}$, $K_{i,t}$, and $L_{i,t}$ denote the output, capital, and labor of firm i. The variable $M_{i,t}$ denotes an aggregate of intermediate inputs consisting of materials sourced from Japan ($m_{i,t}^{J}$) and materials sourced from all places other than Japan ($m_{i,t}^{-J}$), including domestic materials. We are interested in estimating ω and ζ, which parameterize the substitutability between Japanese and non-Japanese materials and that between the capital-labor aggregate and the aggregate of intermediate inputs. The parameters μ_i and ν_i are firm-specific weights and ϕ_i parameterizes the firm's productivity, all of which we assume are constant over the short time horizon we consider. Further, we assume that the firm is monopolistically competitive and faces a CES demand function

$$p_{i,t}^{x} = \left(\frac{Y_{i,t}}{x_{i,t}} \right)^{\frac{1}{\varepsilon}}. \tag{3.8}$$

As usual, $Y_{i,t}$ is the bundle used or consumed downstream and serves as a demand shifter beyond the control of the firm.

3.4.1.1 Pre-Tsunami period

Period τ corresponds to the period April-September 2011, and $\tau-1$ the period September 2010 - February 2011. We exclude the month of March 2011. In period $\tau-1$ the firm operates in a standard environment, choosing capital, labor, and materials so as to maximize

$$p^x_{i,\tau-1}x_{i,\tau-1} - w_{\tau-1}L_{i,\tau-1} - R_{\tau-1}K_{i,\tau-1} - p^{-J}_{i,\tau-1}m^{-J}_{i,\tau-1} - p^{J}_{i,\tau-1}m^{J}_{i,\tau-1}$$

subject to (3.6), (3.7), and (3.8). The firm takes all factor prices as given. Material prices $p^J_{i,\tau-1}$ and $p^{-J}_{i,\tau-1}$ are firm-specific to indicate that different firms use different materials.

It is straightforward to show that this optimization problem implies

$$K_{i,\tau-1} = \frac{\alpha}{1-\alpha}\frac{w_{\tau-1}L_{i,\tau-1}}{R_{\tau-1}}, \tag{3.9}$$

$$\nu_i = \frac{\left(p^{-J}_{i,\tau-1}\right)^{\omega} m^{-J}_{i,\tau-1}}{\left(p^{J}_{i,\tau-1}\right)^{\omega} m^{J}_{i,\tau-1} + \left(p^{-J}_{i,\tau-1}\right)^{\omega} m^{-J}_{i,\tau-1}}, \tag{3.10}$$

$$\mu_i = \frac{\left(\left(\frac{R_{\tau-1}}{\alpha}\right)^{\alpha}\left(\frac{w_{\tau-1}}{1-\alpha}\right)^{1-\alpha}\right)^{\zeta} K^{\alpha}_{i,\tau-1}L^{1-\alpha}_{i,\tau-1}}{\left(P^M_{i,\tau-1}\right)^{\zeta} M_{i,\tau-1} + \left(\left(\frac{R_{\tau-1}}{\alpha}\right)^{\alpha}\left(\frac{w_{\tau-1}}{1-\alpha}\right)^{1-\alpha}\right)^{\zeta} K^{\alpha}_{i,\tau-1}L^{1-\alpha}_{i,\tau-1}}, \tag{3.11}$$

where

$$P^M_{i,\tau-1} = \left[\nu_i\left(p^{-J}_{i,\tau-1}\right)^{1-\omega} + (1-\nu_i)\left(p^{J}_{i,\tau-1}\right)^{1-\omega}\right]^{\frac{1}{1-\omega}}.$$

We will use these relationships in the structural estimation that follows below.

3.4.1.2 Post-Tsunami period

At the beginning of period τ many firms' production processes in Japan are disrupted. Obtaining the desired amount of shipments of materials from Japan may either be prohibitively expensive or simply impossible. Modeling firm behavior in this environment therefore requires modifications to the previous setup. One possibility is to assume that the quantity of materials that firms obtain from Japan is exogenous and that firms freely choose non-Japanese materials, capital and labor. This option is unattractive for two reasons. First, due to existing contracts it is unlikely that a firm is able to adjust the quantities of non-Japanese materials, capital, and labor without costs in such a short time frame. One remedy would be to add adjustment costs to the model. Although straightforward, this approach would require us to estimate additional parameters. Second, and more importantly, the materials sourced from Japan ($m^J_{i,t}$) may not be exogenous for *every* firm. Some suppliers in Japan may have been unaffected by the earthquake and tsunami such that materials could be shipped as desired. Hence, using this approach would require us to distinguish between firms whose supply chains are disrupted and those whose are not. That is, we would have to classify firms based on an endogenous outcome.

For these reasons we prefer an alternative approach, namely to estimate the production function without specifying the full optimization problem. We only assume that in period τ, firms operate the same technologies given by (3.6) and (3.7), and that no firm adjusts its capital stock such that $K_{i,\tau} = K_{i,\tau-1}$. Conditional on knowing the time-invariant features of the production function (ϕ_i, μ_i, ν_i), we next describe an estimation procedure that allows us to find the elasticity parameters most consistent with the observed input choices and output evident in the data.

3.4.2 Estimation

Recall that we use North American exports as a proxy for a firm's output $p^x_{i,t}x_{i,t}$, with the underlying assumption that the former is proportional to the latter. We continue here in the same spirit, though we now make this assumption explicit. Let $V^{NA}_{i,t}$ be the value of North American exports at time t and define

$$\kappa_i = \frac{V^{NA}_{i,\tau-1}}{p^x_{i,\tau-1}x_{i,\tau-1}}. \tag{3.12}$$

In words, κ_i is the fraction of firm i's shipments exported to Canada and Mexico in the six months preceding the tsunami. We next make two assumptions that allow us to construct an estimation equation. First, we assume that a relationship analogous to (3.12) continues to hold in period τ, except for a log-additive error $u_{i,\tau}$. That is,

$$\ln V_{i,\tau}^{NA} = \ln \kappa_i p_{i,\tau}^x x_{i,\tau} + u_{i,\tau}. \tag{3.13}$$

The second assumption is that $\mathbb{E}\left[u_{i,\tau}|X_i\right] = 0$ where X_i is a vector of all right-hand-side variables. Setting the conditional mean of $u_{i,\tau}$ to zero is a standard exogeneity assumption requiring that, loosely speaking, the error is uncorrelated with all right-hand-side variables. It rules out, for example, that in response to a fall in Japanese intermediate imports firms export a fraction of their shipments to Canada and Mexico that systematically differs from κ_i. We provide evidence in section 3.5.3.1 that demonstrates that this is a reasonable assumption.

Using equation (3.6) we can rewrite (3.13) as

$$\ln\left(V_{i,\tau}^{NA}\right) = \ln\left(\kappa_i\phi_i\right) + \ln\left(p_{i,\tau}^x\left[\mu_i^{\frac{1}{\zeta}}\left(K_{i,\tau}^{\alpha}L_{i,\tau}^{1-\alpha}\right)^{\frac{\zeta-1}{\zeta}} + (1-\mu_i)^{\frac{1}{\zeta}}\left(M_{i,\tau}\right)^{\frac{\zeta-1}{\zeta}}\right]^{\frac{\zeta}{\zeta-1}}\right) + u_{i,\tau}. \tag{3.14}$$

Values for ν_i and μ_i are obtained from equations (3.10) and (3.11).[17] Using (3.12), the intercept can be constructed from the previous period

$$\kappa_i\phi_i = \frac{V_{i,\tau-1}^{NA}}{p_{i,\tau-1}^x\left[\mu_i^{\frac{1}{\zeta}}\left(K_{i,\tau-1}^{\alpha}L_{i,\tau-1}^{1-\alpha}\right)^{\frac{\zeta-1}{\zeta}} + (1-\mu_i)^{\frac{1}{\zeta}}\left(M_{i,\tau-1}\right)^{\frac{\zeta-1}{\zeta}}\right]^{\frac{\zeta}{\zeta-1}}}.$$

Notice that κ_i and ϕ_i are not separately identified. Under standard assumptions, we can consistently estimate equation (3.14) using, e.g., nonlinear least squares. The only parameters to calibrate are the rental rate of capital R_τ and the capital share in the capital/labor aggregate α. We estimate the two elasticities, ζ and ω. Notice that ω appears in the intermediate aggregate $M_{i,\tau}$ as shown in equation (3.7). The estimates $(\hat{\zeta}, \hat{\omega})$ solve $\min_{\{\zeta,\omega\}} \sum_{i=1}^{N} (u_{i,\tau})^2$.

[17]After constructing μ_i according to equation (3.11) we average by industry to reduce the level of noise.

Why do we restrict the sample to the year surrounding the Tōhoku event? To understand this, recall that the principal difficulty of estimating production functions lies in unobserved inputs and productivity. Since both are unobserved by the econometrician, they are absorbed into the error term. However, because they are known to the firm, other input choices depend on them. Hence, right-hand-side variables and the error term will generally be correlated, rendering estimates inconsistent.[18]

By restricting the sample period to a single year, the assumption of constant firm productivity seems appropriate. If productivity is constant, it cannot be correlated with the error term, thereby ruling out one of the concerns.[19] The fact that the Tōhoku event was an unexpected shock negates much of the concern about endogeneity arising from unobserved inputs. To see why, consider the case when the firm anticipates a supply chain disruption in a future period. Firm adjustment of unobserved inputs in expectation of this shock will impact input choices – leading to an endogeneity problem where inputs are correlated with the shock. Put simply, the unexpected nature of the Tōhoku event works towards equalizing the information sets between the econometrician and the firm because factor choices are not affected prior to the shock being realized.[20]

Before turning to the data we briefly discuss the intuition of parameter identification. Unlike other approaches to estimating elasticities of substitution (e.g. *Feenstra et al.* (2014)), our method does not rely on the response of relative values to a change in relative prices.[21] In fact, in an econometric sense, our approach treats all inputs as *independent* variables.

A simple example illustrates how the parameters are identified. Consider the production

[18]This problem is discussed in greater detail in, for example, *Ackerberg et al.* (2006).

[19]Of course, the size and exogeneity of the shock also helps with this concern: any idiosyncratic productivity movements during this time are surely subsumed by the earthquake/tsunami.

[20]An unobserved input that could remain operative in our case is that of factor utilization. Since the scope for substantial adjustment along this dimension seems quite limited, we remain confident that our estimates would be robust to the inclusion of this missing ingredient.

[21]Given that we observe little systematic variation in prices (see section 3.5.3.4), we believe that our approach is more appropriate in this setting.

function (3.6) and suppose that, for a particular firm, the initial period yields a value of $(1-\mu) = 0.4$. The elasticity ζ determines how deviations from this measure of the optimal input mix between the intermediate aggregate $M_{i,\tau}$ and the capital labor aggregate translate into measured output. Thus, if we observe comparatively fewer intermediates $M_{i,\tau}$, reflecting a different mix of inputs than that given by 0.4, we obtain an elasticity estimate for ζ that best matches the response in output. Because the estimates for μ, ν, and $\kappa_i\phi_i$ are themselves functions of the elasticities, this procedure must iterate across the parameter space to find the estimate most consistent with the data. Similar reasoning applies for the identification of the ω elasticity based on relative movements in Japanese materials, non-Japanese materials, and output. The estimates we obtain are the best fit across the firms in each sample.

3.4.3 Connecting Model and Data

Estimation of the model requires data on employment, Japanese and non-Japanese material inputs, as well as on exports to North America and output prices for periods $\tau - 1$ and τ. Since data on firm-specific capital stocks are hard to obtain and likely noisy, we use equation (3.9) to construct it from firm payroll and a semi-annual rental rate of 7 percent for period $\tau - 1$.[22] Recall that the capital stock is not adjusted over this time horizon so that $K_{i,\tau} = K_{i,\tau-1}$. The parameter α is calibrated to 1/3.[23]

Quarterly employment information comes from the Business Register, which we adjust to reflect the average value over the 6 month periods we study, as they do not align with the quarters defined within a calendar year.[24] As discussed in earlier sections, the LFTTD contains firm-level data of Japanese imports and North American exports. For non-Japanese

[22]This comes from assuming a real interest rate of 4 percent, combined with an annual depreciation rate of 10 percent, and then adjusting for a semi-annual frequency. The estimates are insensitive to alternative values of the rental rate.

[23]In principle it is possible to construct a firm-specific value for α, using value-added information available in a census year. We are currently exploring the feasibility of this option.

[24]Specifically: $L_{\tau-1} = \frac{1}{6}Emp_{2010Q3} + \frac{1}{2}Emp_{2010Q4} + \frac{1}{3}Emp_{2011Q1}$ and $L_{\tau} = \frac{1}{2}Emp_{2011Q2} + \frac{1}{2}Emp_{2011Q3}$.

material inputs, we would ideally combine the non-Japanese imported materials with information on domestic material usage for these firms. As information on domestic material inputs is not available in Census data at this frequency, we utilize information on the *total* material expenditures from the Census of Manufacturers (CM) to construct a firm-level scaling factor to gross up non-Japanese intermediate imports. Put differently, we impute non-Japanese material inputs from non-Japanese input imports. For each firm, we construct the scaling factor as

$$\frac{P_i^M M_i - p_i^J m_i^J}{p_i^{-J} m_i^{-J}} \tag{3.15}$$

from the latest CM year. Because the closest available CM year is 2007 in our data, there is some concern about missing or outdated information for this factor. We mitigate this by using industry-specific means for missing values, and winsorizing large outliers at the 90th/10th percentiles.

Regarding information on prices, the LFTTD records the value and quantity of each trade transaction (at the HS10 level), and thus it is possible to construct the associated price, or "unit-value" of each shipment directly.[25] Aggregating up these shipments into a firm-month observation is complicated, of course, by the differing quantity units. Lacking any better alternative, we simply average the transaction prices using the dollar value of each transaction as weights.

Finally, we restrict the sample of firms to those that have regular imports from Japan and non-Japan over the periods we study, as well as regular North American exports.[26] While this substantially limits the number of firms in each sample, the shares of trade represented by these firms in each category remains very high (see Table 3.3).

We obtain standard errors using bootstrap methods, which also allow us to account for the

[25] Those transactions with missing or imputed quantity information are dropped. Future efforts will evaluate whether it is possible to recover the quantity values from prior transaction details.

[26] Specifically, we drop any firm that has more than 3 months of zeros for any of these values, over the period $\tau - 1$ or the period τ.

uncertainty implied by the imputation of non-Japanese material inputs. We draw randomly with replacement from our set of firms to construct 5000 different bootstrap samples. For each of these samples, the non-Japanese materials share is imputed as described above before the estimation proceeds.

3.4.4 Summary of Results

The results of the estimation are shown in Table 3.3. The elasticity between material inputs for Japanese affiliates is 0.2, while the elasticity between the aggregate material input and capital/labor is 0.03. Together, these estimates are indeed consistent with the reduced-form evidence for the (ψ) elasticity from section 3.3.2. The relative magnitudes are also intuitive: while Japanese imported inputs are strong complements with other material inputs — consistent with the high share of intra-firm transactions comprising this trade — there is even less scope for substitution between material inputs and domestic capital/labor.

The estimation procedure also allows us to estimate these elasticities for two samples of non-Japanese firms: non-Japanese multinationals and non-multinational firms. While the estimates for the ζ elasticity are indeed very close for these other samples, the elasticity estimates corresponding to material inputs are higher, at 0.6 and 0.4 respectively. The lower share of intra-firm imports from Japan for the non-Japanese multinationals aligns with the argument that this type of trade is the key source of non-substitutability in the short-run. On the other hand, the low estimates for non-multinational firms, which have essentially zero intra-firm imports, may point to other mechanisms at work beyond the role of intra-firm trade. More generally, however, the estimates for these parameters are all significantly lower than what is commonly assumed (typically unity or higher) in the literature.

Although the number of firms included in this estimation is small (550 firms in total across the three subgroups) , they account for a large share of economic activity in the United States. Looking at their combined share of total trade, these firms account for over

80% of Japanese intermediate imports, 68% of non-Japanese intermediate imports, and well over 50% of North American exports. Such high concentration of trade among relatively few firms is consistent with other studies using this data (see *Bernard et al.* (2007)).

3.5 Discussion

The structural estimates of the model are broadly in agreement with the evidence in section 3.3.2: imported inputs are strong complements with other inputs in the production function. The rigidity of the production function for multinational firms in particular is likely due to i) the high degree of intra-firm trade in what is presumably highly specialized inputs, and ii) .[27] Our results have a number of important implications for how economists should think about multinational firms in general, as well as aggregate topics such as volatility and business cycle co-movement.

3.5.1 Aggregation

Before relating our estimates to macroeconomic topics, it is important to discuss aggregation. Indeed, in any study utilizing micro-level estimates to inform macro-level objects of interest, the details of aggregation and heterogeneity are of critical importance. Work by *Imbs and Méjean* (2011) argues that imposing homogeneity across sectors when estimating consumption elasticities can be overly restrictive, creating a heterogeneity bias which can be quantitatively large. In our case one could discuss aggregation along various dimensions: across products, industries, firms, and so on. We examine the effects of product-level aggregation in section 3.5.3.3 below.

A primary concern is how to translate the results from the firm-level subsamples into estimates that would pertain to macro-oriented models. As a first step, the final column

[27] The vertical integration of production across countries, within the firm, has shown to be a key driver of the decline in joint ventures (see *Desai et al.* (2004)).

in Table 3.3 shows the elasticity estimates when aggregating across all firms in the sample. The results are consistent with the estimates by subgroup, suggesting substantial complementarities across inputs. All estimates in Table 3.3, however, correspond to the average across firms in each group, and do not take into account heterogeneity in firm size within the groups. It is relatively straightforward to modify our estimation procedure to weight firms according to their relative size.[28] We report the results from this modified estimation in Panel B of Table 3.4. When comparing the results to those in Table 3.3, it is evident that the weighted estimates are not substantially different than the unweighted estimates. Although the samples of firms comprising these estimates do not amount to the total manufacturing sector of the United States, they do account for the considerable majority of U.S. trade.

3.5.2 Implications

The rigid production networks of foreign-owned multinationals will have direct consequences on the destination (host) economy. Previous literature has hypothesized that input linkages could generate business-cycle comovement, but supportive empirical evidence has been difficult to find. This paper can be seen as a first step in establishing empirical evidence for a causal relationship between trade, multinational firms, and business cycle comovement. In a companion paper (*Boehm et al.* (2014)), we evaluate the quantitative importance of such complementarities of imported inputs by multinational affiliates. When separately accounting for intermediate input trade by multinationals and traditional trade in final goods, the model distinguishes between the production elasticity of imported inputs and the traditional "Armington" elasticity used to bundle together international goods for consumption. The complementarities in import linkages by multinationals increases value-added comovement

[28]Since the appropriate measure of size in our context is output, we follow our convention and use the relative amounts of North American exports in the period before the shock as the weights.

in the model by 11 percentage points relative to a benchmark without such firms.

This model shares similarities with several other existing models, particularly *Burstein et al.* (2008). A key advantage of *Boehm et al.* (2014), however, is a tight link to Census data for matching other features of multinationals and trade. *Johnson* (2014) also looks at the role of vertical linkages on comovement, but applies greater input-output structure on the model. Such features will generate increases in value-added comovement in his model, the magnitude of which becomes significant only when the elasticity of substitution among inputs is sufficiently low. Other work also identifies multinationals as a key source of the transmission of shocks: *Cravino and Levchenko* (2014) demonstrates that foreign multinational affiliates can account for about 10 percent of aggregate productivity shocks.[29]

The low value for ω indicates the presence of spillovers beyond the immediate effect from Japan. That is, imports from non-Japanese locations are lower as a result of the shock in Japan[30], and we would presume this applies to suppliers within the United States as well. Specifically, upstream suppliers (in countries other than Japan as well as within the U.S.) were affected indirectly due to the exposure of Japanese affiliates to the shock combined with the rigidity of their production with respect to those inputs. Downstream suppliers that rely on the inputs from the disrupted firms would likewise be adversely affected. The presence of such spillovers combined with the large network of input linkages can indeed magnify the total effect of the transmission of the shock to the U.S. market. Such effects are also evident in a related paper, *Carvalho et al.* (2014), which finds large spillovers in both upstream and downstream firms in Japan following the 2011 earthquake.

Another branch of literature on the diversification of risk has studied whether firms using complex production structures with several intermediates could be less volatile (*Koren and*

[29]Of course, shocks can be passed through to affiliates through other means as well. See *Peek and Rosengren* (1997) and *Peek and Rosengren* (2000) for the case of U.S. affiliates from Japan.

[30]To confirm this, see Figure **??**, which replicates the results in Figures 3.6 (Panel A) and 3.7, but only for non-Japanese imports.

Tenreyro (2013)). *Kurz and Senses* (2013) establish that firms with substantial imports and exports have lower employment volatility than domestic firms in the medium to long term, which they attribute partly to the diversification of risk.[31] The key result in this paper points to a possibly overlooked fact: the extent of the benefits from diversification depends heavily on the substitutability of inputs. Conditional on a given number of inputs used in production, a firm will likely experience greater volatility if each input is key to the production process and inputs are subject to heterogeneous shocks.[32] Conceptually, an increase in the use of imported inputs should not be viewed necessarily as diversification. A fragmentation of production can lead to an increased supply chain risk that is an important counterweight to whatever efficiencies such complex input sourcing might afford, particularly when the production elasticities are low.

The rigid production networks of multinational firms also influences our understanding of why firms segment production across country borders. In a related paper, *Flaaen* (2013a) shows that despite the presence of substantial and complex import linkages with the source country (consistent with a vertical framework of FDI), the motive for multinational production appears to be to serve the domestic market (consistent with the horizontal framework of FDI). The result could be called "horizontal FDI with production sharing."[33] The evidence for strong complementarities in this production sharing, however, presents a puzzle. Why does the firm replicate only select portions of the supply chain, considering the penalties for disruptions and mismatched inputs are so great? It is perhaps the case that the segments of the production chain that remain in the source country have a location-specific component

[31] An interesting result from *Kurz and Senses* (2013) is that firms that only import are actually more volatile than the domestic-only benchmark.

[32] *Pravin and Levchenko* (2014) outline theoretical results showing that for a given elasticity value (in their case, Leontief), volatility in output per worker should be actually decreasing in the number of inputs used.

[33] *Ramondo et al.* (2014) is another recent example arguing for a more nuanced interpretation of multinational production.

that is not easily transferable when the firm moves production abroad.[34] Understanding the dynamics behind these sourcing decisions is an area in need of further research.[35]

3.5.3 Robustness and Extensions

3.5.3.1 Mis-measurement of Firm Production

A natural concern with our analysis is the use of N.A. exports as a proxy for firm-level production. Perhaps it is the case that shipments abroad fall disproportionately more than domestic shipments following a shock to production. If this were the case, the N.A. exports would indeed be a poor proxy for production, and its usefulness in evaluating a production elasticity substantially compromised.

To evaluate this concern, we narrow our study to the automotive sector, which has data on production, sales, and inventory at a monthly frequency. Using the Ward's electronic databank, which reproduces the published series in the annual Automotive Yearbook, we obtain plant-level information on production, and model-line information on inventory, sales, and incentives.[36] The baseline specification is the same as in equation (3.5), where the dependent variable is now Q_{jit}: production of plant j of firm i in month t. The Japanese multinational firms are, in this case, those automakers with plants located within North America but whose parent company is headquartered in Japan.[37]

Figure 3.8 shows the results, where we once again divide by pre-shock levels to gain a sense of the percentage effects of these changes. Relative to their U.S. counterparts, Japanese automakers in the United States experienced large drops in production following the Tōhoku

[34] The model of knowledge sharing in *Keller and Yeaple* (2013) is one attempt to analyze the dynamics between such transfers being accomplished in embodied (intra-firm trade) or disembodied (direct communication) form. Alternatively, domestic content requirements may provide incentives to produce specified inputs in one location over another.

[35] For a recent example of how such investment and sourcing decisions can alter a country's comparative advantage over time, see *Alviarez* (2014).

[36] Appendix C.3.7 details further features of this data and explains how the sample was constructed.

[37] These firms are Honda, Mitsubishi, Nissan, Toyota, and Subaru.

event. Production bottomed out in May of 2011 — two months after the event — at almost 60 percent below trend. [38] The point estimates return to a level near zero in September of 2011, implying that the shock affected production for nearly 6 months.[39] We interpret these results to be largely supportive of the results obtained using the exports-based proxy for production. The percentage drops in the two series are remarkably similar: a trough of 59% at $t = 2$ in the automotive data vs 53% at $t = 2$ using the proxy. We conclude that, at least for this exercise, the proxy appears to be providing valuable information on a firm's U.S. production behavior.[40]

3.5.3.2 Intermediate Input Inventories

Inventories are another obvious feature that should influence the relationship between input shipments, production, and the elasticity of substitution. In particular, inventories of intermediate inputs allow the firm to absorb unforeseen shocks to input deliveries without an impact on the production process.[41] As it relates to the production elasticity, however, the presence of these inventories should serve to diminish or delay the production impact, thereby *increasing* the elasticity relative to what it would be without such inventories.

In fact, it is striking the extent to which we do not see any evidence for the role of

[38]The average monthly plant-level production at these firms during December 2010 through February 2011 was about 12,200 units a month. The magnitude of the drop in May was -7200 units.

[39]We describe additional results on the behavior of inventories, sales, incentives, and production in Japan in an appendix.

[40]In addition, one might be concerned that the N.A. exports series may be contaminated with Japanese imports whose country of ultimate destination is Canada/Mexico (a.k.a "in-transit shipments" – imports to Canada/Mexico via U.S.). These shipments should not be picked up in the reporting systems underlying the LFTTD. According to section 30.2(d)(1) of the U.S. Code of Federal Regulations, "In-transit shipments of goods from one foreign country to another where such goods do not enter the consumption channels of the United States are excluded from filing the Electronic Export Information (EEI)." Additionally, the Army Corps of Engineers has suspended the requirement to file the Form 7513, Shippers Export Declaration (SED) for In-transit Goods leaving the United States via vessel. Finally, the corroborating results from section 3.5.3.1 should also serve to allay such concerns.

[41]The existence of final good inventories, on the other hand, makes a distinction between the production and sales of a particular product. Here, the presence of final good inventories implies that the firm can continue to sell from existing inventory stocks even while production is temporarily affected.

intermediate input inventories in the production impacts of Figure 3.6 (Panel B) or Figure 3.8. The effect on production appears to be almost immediate, indicating that the stock of inventories of imported intermediates is low (less than one month's supply) for these firms.

We obtain a rough sense of the degree of inventory holdings from the Census of Manufacturers micro-data. Combining information on the beginning period stock of materials inventories with the annual usage of materials, we calculate the average monthly supply of inventories for each firm.[42] Panel A of Table 3.1 calculates the production-weighted averages over a select set of firm groups.[43] We see that on average, Japanese multinationals hold a little over 3-weeks supply of intermediate inputs as inventory. This is slightly less than non-multinational firms, a fact that aligns with the oft-cited "lean" production processes made famous by Japanese firms in previous decades. Though these data are for the year 2007, there is little reason to believe these relative magnitudes have changed substantially over a period of a few years. For completeness, Panel A of Table 3.1 also reports the corresponding estimates for output inventories.[44]

Low inventory holdings combined with an inelastic production function suggests that firms are willing to tolerate some degree of expected volatility in their production. Either the costs of holding inventories or diversifying sources of supply are sufficiently high, or firms believe the probability of disruption is low. In either case, these lean production strategies carry a greater potential for the propagation of shocks across countries, perhaps affecting firms with limited knowledge of their indirect exposure through complicated production chains.

[42] Unfortunately, the CM data does not report imported materials inventory separately.

[43] These numbers are broadly similar, though somewhat lower than other estimates in the literature. See *Ramey* (1989) for one example.

[44] At first glance, the average monthly supply of these output inventories looks surprisingly low. On the other hand, it is probably the case that inventories are held jointly by the manufacturer and wholesale/retail establishments Thus, considering the inventories of manufacturers alone could potentially under-represent the "true" level of output inventories available for smoothing out production disturbances.

3.5.3.3 Multi-Products and Sub-Optimal Mix

In the frameworks used in sections 3.3.1 and 3.4, we consider the aggregate bundles of imported intermediates, abstracting away from product-level detail. In reality, the firms in our dataset often import many distinct intermediate inputs from Japan. The structure of a CES production function implies that if each of these within-country inputs was non-substitutable with one another (a further, nested Leontief structure), the production impact of a disruption in the supply of just one input could be amplified relative to the value of that input.[45] We evaluate this possibility below.

This is particularly true given the heterogeneous impact of the Tōhoku event across Japan (see Figure 3.4). This could translate into considerable dispersion in the impact on the *products* imported by a particular U.S. firm or Japanese affiliate. With product-level shocks, considering the effect on the aggregate import bundle amounts to assuming either 1) perfect substitutability among products, or 2) that the firm maintains an optimal within-country product mix at all times.

To be concrete, it may be more accurate to view the M_t in equation (3.1) as a further C.E.S. function of multiple products. Thus, we can define the proper measurement of this variable as

$$V_{i,t}^{M} = P_{i,t}^{M} \left(\sum_{n=1}^{N} \eta_n^{\frac{1}{\chi}} (m_{n,i,t}^{J})^{\frac{\chi-1}{\chi}} \right)^{\frac{\chi}{\chi-1}}, \tag{3.16}$$

where now $V_{i,t}^{M}$ is the value based on a combination of N distinct products, with weights η_n and elasticity χ.

Product-level heterogeneity in the production impact of the shock combined with imperfect coordination among input suppliers implies that the aggregate (measured) import bundle for a particular firm may turn out to be suboptimal. In this case, we are measuring $\widehat{V^M}_{i,t} = \sum_{n=1}^{N} (p_{n,t}^{m} m_{n,t}) \geq V_{i,t}^{M}$. And the lower the elasticity of substitution among products,

[45] This point has been made in somewhat differing contexts, by *Kremer* (1993) and *Jones* (2011).

the more severe the disconnect between the measured imports and the "effective" imports — that which is actually useful in downstream production.

A suboptimal product mix indicates that measured imports ($\widehat{V^M}_{i,t}$) are greater than the effective imports ($V^M_{i,t}$). As a result the measured output response to the import shock will be larger than otherwise, resulting in a downward "bias" in the elasticity estimates from section 3.3.1 and 3.4.[46] Such an effect is decreasing in the product-level elasticity parameter χ, as complementarity itself is the driving force between differences in $\widehat{V^M}_{i,t}$ and $V^M_{i,t}$. In addition, the effect is also increasing in the degree of deviation from the optimal product mix.

Does this exert a quantitatively large effect on our point estimates? Given the emphasis on low inventories and lean production processes in downstream operations, one might expect that across-product adjustment would take place before sending the inputs abroad. To analyze this empirically, we analyze whether there are significant deviations in the product composition of Japanese imports during the months following the Tōhoku event. To do this, we construct a measure of the distance of a firm's import bundle from a benchmark, which we will interpret to be the optimal bundle. Let $t = s^*$ be such a benchmark date. Then, using the product-level information in the LFTTD data we construct for each firm i, the share of total imports from Japan for a given product code n. Defining this share to be $s_{n,i,t}$, we then construct the average product-level distance from optimum $DO_{i,t}$ as:

$$DO_{i,t} = \frac{1}{N^i} \sum_{n=1}^{N^i} \left(|s_{n,i,t} - s_{n,i,s^*}| \right) \tag{3.17}$$

where N^i is the total number of products imported by firm i. We define the period s^* to be the months of April-June of 2010, and then evaluate DO_i at a monthly frequency, with particular interest in the months following the Tōhoku event. While there may be

[46]Because the source of this downward pressure on the estimate for ψ (or ω) is itself a very low product-level elasticity, it is unclear whether this should be considered a bias in the traditional sense.

natural movements in the bundle of products imported from Japan, evidence for substantial coordination failure in product composition or heterogeneity in product-level shocks would come from any abnormal jumps in this index in the months of the disruption. One can calculate this at various levels of product aggregation (i.e. HS4, HS6, HS8, HS10), though we report results using the HS6 level.[47]

The results of this exercise are shown in Figure 3.9. We plot the average DO_i across Japanese firms for each month (the figure shows a 3-month moving average) during the period 2009-2011. Mechanically, this measure should be relatively close to zero in the months consisting of the benchmark (April-June 2010). While there is a secular rise in this measure on either side of this benchmark period, there do not appear to be any large jumps in the months directly following the Tōhoku event. More interesting, perhaps, are the considerably larger values for this measure during early 2009, which might reflect the effects of the trade collapse associated with the Great Recession. We interpret Figure 3.9 as evidence that the potential for suboptimal mix across products from Japan does not pose a serious problem to our measurement in previous sections.

3.5.3.4 Other Considerations

Strategic Behavior: Another possibility that could affect the interpretation of the results from Figure 3.7 might be strategic behavior, particularly on the part of the competitors of Japanese firms in the United States. These firms could raise production or prices following the negative supply shock affecting their competitors, which would serve to bias downward the β_p coefficients from the equation with $X_{i,t}^{NA}$ as the dependent variable.[48] To evaluate this

[47]The level of aggregation we use attempts to balance concerns along two dimensions. With less product aggregation (i.e. HS10 level), one might be concerned with the inherent lumpiness of product-level firm imports. More product aggregation, on the other hand, could mask important product differences within a particular product grouping.

[48]Specifically, in equation (3.5) the γ_p's would be higher than would be expected without the shock, and hence the β_p's artificially low.

possibility, we turn once again to the automotive data. Here, we can look directly at the production of non-Japanese automakers in the months directly following the Tōhoku event. Appendix Figure A1 plots the relative production of these firms, using time-series variation only. There appears to be no quantitatively meaningful responses in the months following March 2011. This should not come as a surprise given capacity constraints and utilization adjustment costs, particularly given the short time horizon. We provide evidence on the role of prices next.

Prices: Traditionally, estimating the elasticity of substitution is accomplished via price and quantity data for products over extended periods of time. For the short horizon we consider in this paper, there are several reasons why prices may not have the scope to adjust. Many supplier relationships negotiate prices for longer periods of time than one or two months. Second, and perhaps more importantly, Table 3.2 demonstrates that the large majority of imported intermediate inputs are intra-firm. The observed prices of these transactions are transfer-prices (within firm) and not likely to change reflecting any short-term disturbance. However, because the LFTTD contains both quantity and price information, we can confirm whether or not prices remained relatively stable during this period. The results in Appendix Table C.6 confirm that there are few significant price movements on import or export transactions for either Japanese or non-Japanese multinationals surrounding the Tōhoku event.[49]

Domestic Inputs: It is also possible to evaluate the response of domestic inputs directly, using the limited information we have on quarterly firm-level employment and payroll information, taken from the Census Bureau's Business Register (BR).[50] We consider the evidence in Appendix C.3.5 and find no significant effects on either employment or payroll

[49]Further details on the construction of the data underlying the analysis of unit values is available in Appendix C.3.6.

[50]The BR itself receives quarterly payroll and employment information for business and organizational employers from the IRS: Form 941, the Employer's Quarterly Federal Tax Return. For more information on the BR (formerly the SSEL), see *Walker* (1997).

for Japanese firms in the quarter(s) following the shock (see Table C.5). Of course, there are a number of reasons — principally labor adjustment costs — why one would expect little, if any, impacts on employment following this short-lived shock. Press releases dispatched by the Japanese automakers during this time indicated that no layoffs would occur. Rather, the firms indicated that they would use the production stoppages for employee skill and safety training.

3.5.4 External Validity

Finally, we discuss the external validity of this result. The exogenous variation we use to identify this elasticity is tied to a particular event in time, making generalization subject to some caveats. On the other hand, there are few, if any, estimates of this parameter in the existing literature. The critical question is whether the mechanisms underlying the elasticity estimates are operative beyond the circumstances surrounding this event study.

The pattern of strong intermediate input linkages with the source country is not restricted to Japanese affiliates only. As shown in *Flaaen* (2013b), over 45 percent of the imports for all foreign multinational affiliates are sourced from the country of the parent firm. The cost share of imported intermediates from the source country is 0.12 for all foreign affiliates, which is lower than the 0.22 for Japanese affiliates but still much larger than the representative importing firm in the United States. The cost share of all imported inputs is actually quite close: 35 percent for Japanese affiliates vs 32 percent for all foreign affiliates.

A related concern is whether the estimates for Japanese affiliates are driven solely by the automotive sector. The ideal check would be to run industry-by-industry subgroup estimates for the elasticities, thereby generating heterogeneity that could be assessed relative to expectations. Unfortunately, the small number of firms applicable for this analysis, combined with disclosure requirements associated with the Census Bureau data usage, prevents this degree of detail. Instead, we address this concern by splitting the sample into a motor vehicle and non-motor vehicle subsample. We do this for the Japanese multinationals as well as the total sample of all firms. The results for these four subsamples are reported in Panel C of Table 3.4. Using the published data from the B.E.A., the automotive sector is a large but not overwhelming percentage of total Japanese manufacturing affiliates in the U.S. The entire motor vehicle sector as a whole comprises significantly less than half of value-added (roughly 40 percent) for the Japanese manufacturing affiliates.

When viewed in light of the substantial fraction of intra-firm imports comprising multinational affiliate trade, the low elasticity of substitution should not come as a surprise. One would not expect close substitutes for the sort of specialized products reflecting firm-specific knowledge that likely comprises this trade. Moreover, such a low estimate for an elasticity of this nature is not without precedent. Using different methodologies, recent work by *Atalay* (2014) highlights strong complementarities between intermediate inputs, using industry-level data for the United States.[51]

Any elasticity estimate is tied to the time-horizon to which it corresponds. *Ruhl* (2004) emphasizes the difference between elasticities implied by responses to temporary vs permanent shocks. Larger values are calculated for an elasticity following a permanent shock, owing in part to firm responses along the extensive margin. In our context, we estimate the elasticity subject to a short-lived shock where the structure of the supply chain is plausibly fixed and extensive margin movements of supplier relationships would not apply. For this reason the elasticity parameters (ω, ζ) should likely generalize to other contexts of this horizon and for shocks of this general duration. Even for a long-lived shock, the estimated elasticities would remain relevant while the firm makes changes to its network of suppliers. Evaluating whether there is evidence for long-term supply-chain reorganization following the Tōhoku event is an area of ongoing work.

[51] The point estimate for the elasticity of substitution among intermediate inputs from *Atalay* (2014) is 0.03.

3.6 Conclusions

Using a novel firm-level dataset to analyze firm behavior surrounding a large exogenous shock, this paper reveals the mechanisms underlying cross-country spillovers. We find complementarities in the international production networks of Japanese affiliates, such that the U.S. output of these firms declined dramatically following the Tōhoku earthquake, roughly in line with an equally large decline in imported inputs. The elasticity of substitution between imported and domestic inputs that would best match this behavior is very low – nearly that implied by a Leontief production function. The reliance on intra-firm imports by multinational affiliates from their source country is the most plausible explanation for such strong complementarities in production. Structural estimates of disaggregated elasticities are similarly low, and imply spillovers to upstream and downstream firms in the U.S. and abroad (non-Japan). The large impacts to Japanese affiliates together with the propagation to other U.S. firms explains the large transmission of the shock to the U.S. economy in the aggregate.

These elasticities play a critical role in the way international trade impacts both source and destination economies. Such complementarities between domestic and foreign goods have been shown to improve the ability of leading theoretical models to fit key moments of the data. We emphasize here the distinction between substitutability between domestic and foreign final goods (a "consumption" elasticity of substitution, or the so-called Armington elasticity) and substitutability between domestic and foreign intermediate goods (a "production" elasticity of substitution). In a companion paper (*Boehm et al.* (2014)), we document the behavior of a model with such complementarities in imported intermediates, and discuss how these elasticity parameters interact. Calibrating this model to the share of multinational affiliate trade in intermediates yields an increase in value-added comovement of 11 p.p.

Such rigid production networks will also play a substantial role in aggregate volatility, productivity growth and dispersion, and the international ownership structure of production. The novel datasets described in this paper may help to shed light on these and other areas of research in the future.

Table 3.1: Summary Statistics: Imported Inputs and Inventories by Firm Type

	Japanese Multinationals	Non Multinationals
Panel A: Avg. Monthly Supply of Inventories		
Inputs	0.83	1.08
Output	0.31	0.45
Panel B: Cost Share Of Imported Inputs		
from Japan	21.8	1.0
from all countries	35.0	17.5

Source: CM, LFTTD, DCA, and UBP as explained in the text. The data are for year 2007. This table reports the average monthly supply of inventories [(usage/12)/beginning period inventory stock] for materials and output, as well as the cost share of imported products.

Table 3.2: Summary Statistics

Panel A: Cost Share Of Imported Inputs

	Japanese Firms	Non Multinationals
from Japan	21.8	1.0
from all countries	35.0	17.5

Panel B: Treatment Effects Sample Details

	Japanese Firms	Other Multinationals	Balancing Tests t	$p > \|t\|$	% Reduct |bias|
N.A. Exports	3,504,894	3,413,058	0.38	0.706	79.1
share intra-firm	72.0	52.2			
Intermediate Input Imports	8,075,893	7,596,761	0.87	0.384	88
share from Japan	70.0	3.5			
share intra-firm	86.0	21.7			
Industry (Avg)	–	–	0.009	0.965	97.8

Source: LFTTD, DCA, and UBP as explained in the text.
Panel A data are for year 2007. Panel B reports the baseline average values of N.A. exports and intermediate input imports, as well as the characteristics of that trade, for the two groups of firms: Japanese affiliates and other multinational firms. The statistics are calculated in the three months prior to the Tōhoku earthquake: Dec. 2010, Jan. 2011, and Feb 2011. The control group of other multinational firms has been re-weighted using the normalized propensity score, from a specification including the level of N.A. exports, int imports, and industry dummies. The final three columns report balancing tests of the equality of the means between the treated and control group.

Table 3.3: Firm-Level Estimation: Results and Sample Details

Panel A: Calibration

Parameter	Value
R_t	0.07
α	1/3

Panel B: Estimation Results

	Japanese Multinationals	Non-Japanese Multinationals	Non-Multinationals	All Firms
ω	0.201	0.624	0.423	0.552
	[0.02 0.43]	[0.16 0.69]	[0.26 0.58]	[0.21 0.62]
ζ	0.032	0.038	0.032	0.037
	[0.030 0.673]	[0.035 0.508]	[0.029 1.68]	[0.034 0.038]
	Sample Details			
Weight on K/L Aggregate ($\bar{\mu}$)	0.223	0.514	0.278	0.409
Weight on JPN Materials ($1-\bar{\nu}$)	0.173	0.044	0.147	0.096
Number of Firms	105	304	141	550
Share of Total Trade				
JPN int imports	0.60	0.23	0.03	0.86
Non-JPN int imports	0.02	0.66	0.01	0.69
N.A. exports	0.08	0.47	0.01	0.56

Source: CM, LFTTD, DCA, and UBP as explained in the text.
This table reports the results from the firm-level estimation detailed in section 3.4. Panel A outlines the parameters that are calibrated prior to estimation. The top two rows of Panel B reports the point estimates of the elasticities, and the corresponding 95 percent confidence intervals using a bootstrapping procedure. (See Appendix C.3.3 for more details on the measurement of dispersion for these estimates.) Rows 3 and 4 report other estimates related to the calculated production functions. The final rows of Panel B describe features of the estimation samples.

Table 3.4: Firm-Level Estimation: Other Results

Panel A: Calibration

Parameter	Value
R_t	0.07
α	1/3

Panel B: Estimation Results (Weighted)

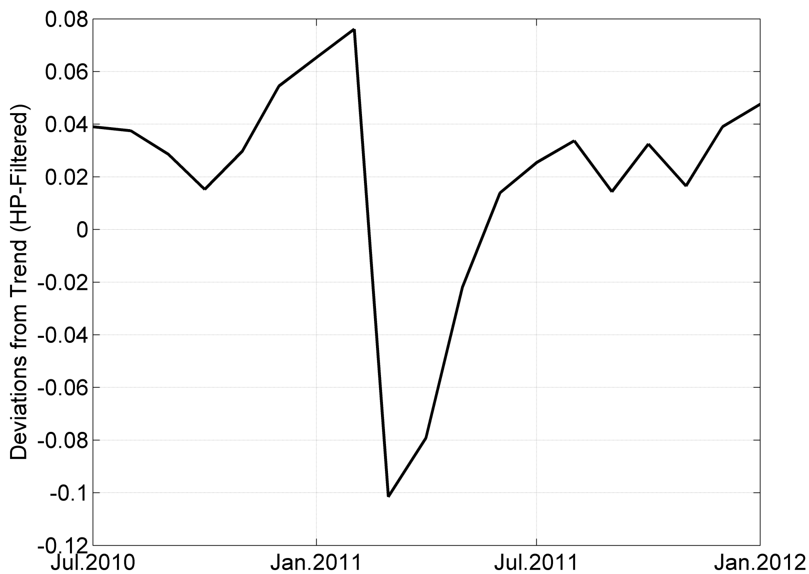
	Japanese Multinationals	Non-Japanese Multinationals	Non-Multinationals	All Firms
ω	0.157	0.611	0.543	0.606
	[0.02 0.40]	[0.30 1.23]	[0.305 0.57]	[0.28 0.70]
ζ	0.241	0.038	0.032	0.037
	[0.03 0.884]	[0.034 0.51]	[0.029 0.55]	[0.034 0.038]
Number of Firms	105	304	141	550

Panel C: Estimation Results: MV Sector

	Japanese Mult.		All Firms	
	Motor Vehicles	Non-Motor Vehicles	Motor Vehicles	Non-Motor Vehicles
ω	0.311	0.094	0.414	0.574
	[0.019 0.398]	[0.016 0.59]	[0.27 0.60]	[0.16 0.66]
ζ	0.032	0.071	0.037	0.037
	[0.030 0.48]	[0.028 1.27]	[0.031 0.64]	[0.033 0.037]
Number of Firms	35	70	100	450

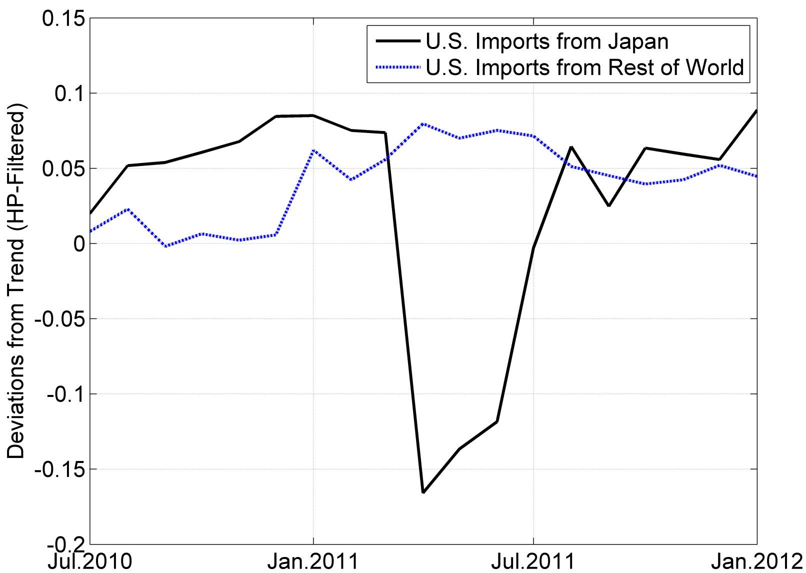
Source: CM, LFTTD, DCA, and UBP as explained in the text.
This table reports additional estimation results. Panel B recalculates the results from Table 3.3 using a vector of weights to assign larger firms a greater share in the estimation. Panel C divides the samples based on the motor vehicle industry.

Figure 3.1: Index of Japanese Industrial Production: Manufacturing Jul.2010 - Jan.2012



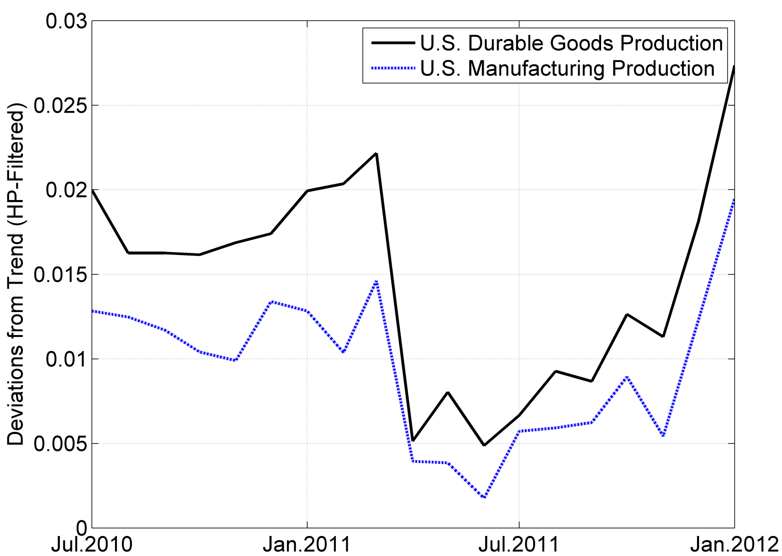
Source: Japanese Ministry of Economy, Trade, and Industry (METI). The series are logged, HP-Filtered, after seasonally adjusting.

Figure 3.2: U.S. Imports from Japan and Rest of World, Jul.2010 - Jan.2012



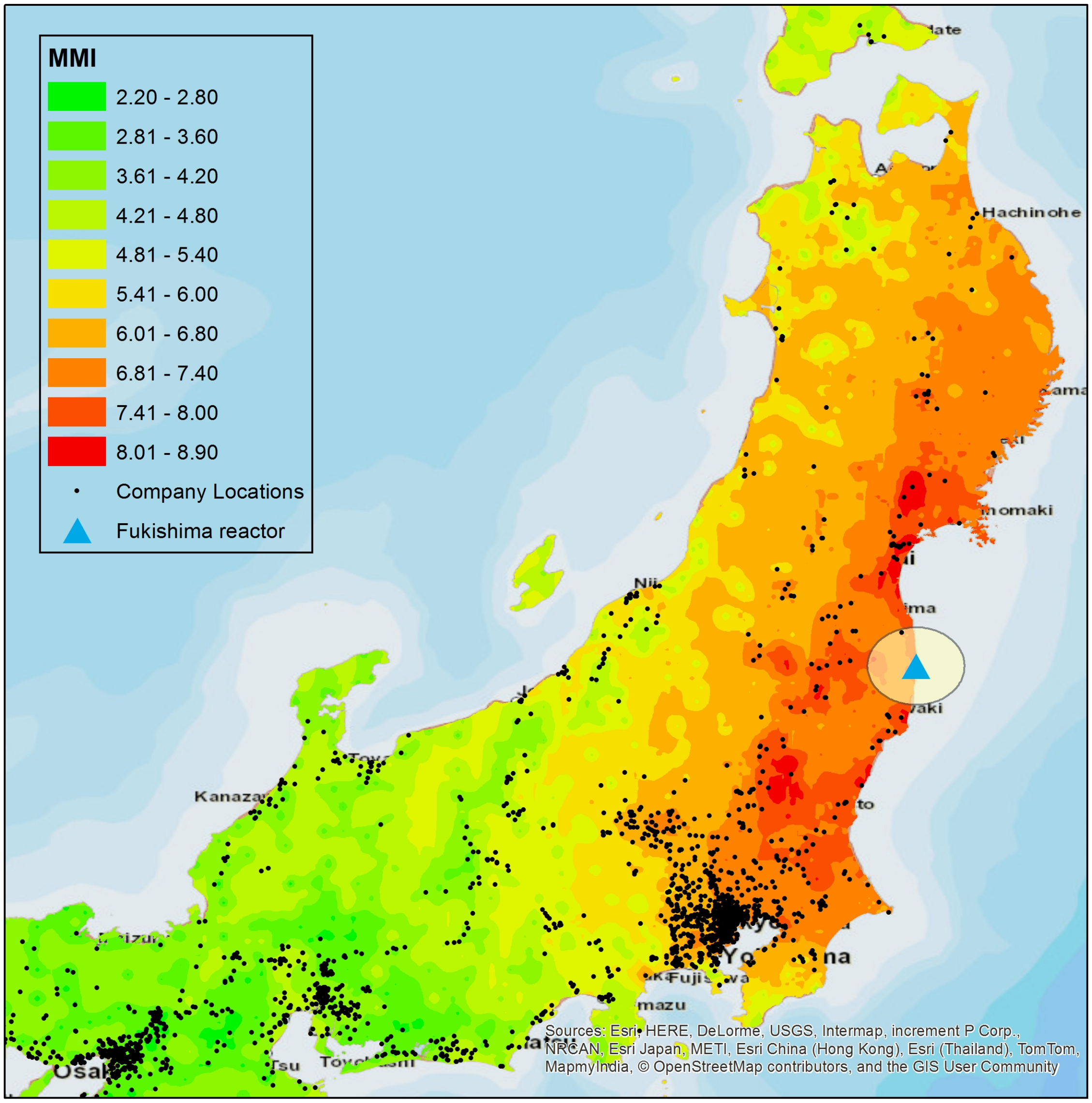
Source: U.S. Census Bureau (FT900: U.S. International Trade in Goods and Services). The series are logged, HP-Filtered, after seasonally adjusting.

Figure 3.3: U.S. Industrial Production: Manufacturing and Durable Goods



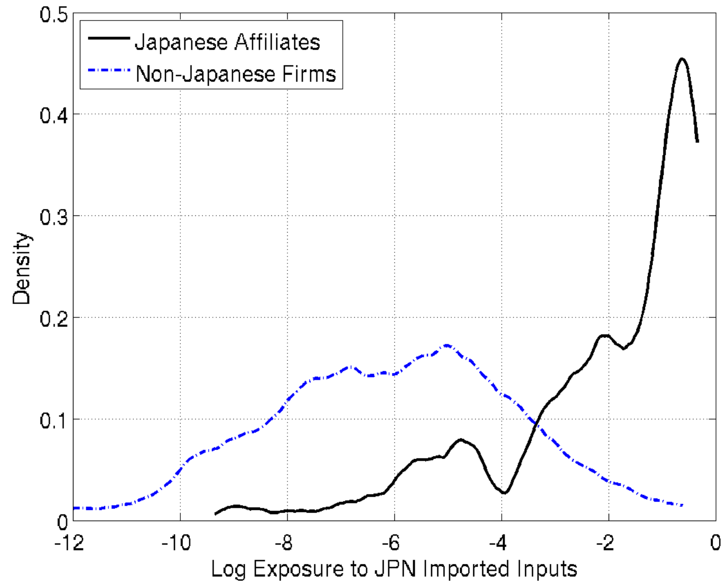
Source: Federal Reserve Board, Industrial Production and Capacity Utilization - G.17 Series . Series is Seasonally Adjusted.

Figure 3.4: Geographic Distribution of Earthquake Intensity and Affiliate Locations



Source: USGS and DCA/Uniworld Directories
This figure plots the geographic distribution of the Tōhoku earthquake, based on recorded measurements taken directly after the event. The "Modified Mercalli Intensity" (MMI) scale is constructed based on a relation of survey response and measured peak acceleration and velocity amplitudes from prior major seismic events. Each dot corresponds to a geocoded Japanese affiliate location corresponding to a firm with U.S. operations. For more details, see Appendix C.2.3.2.

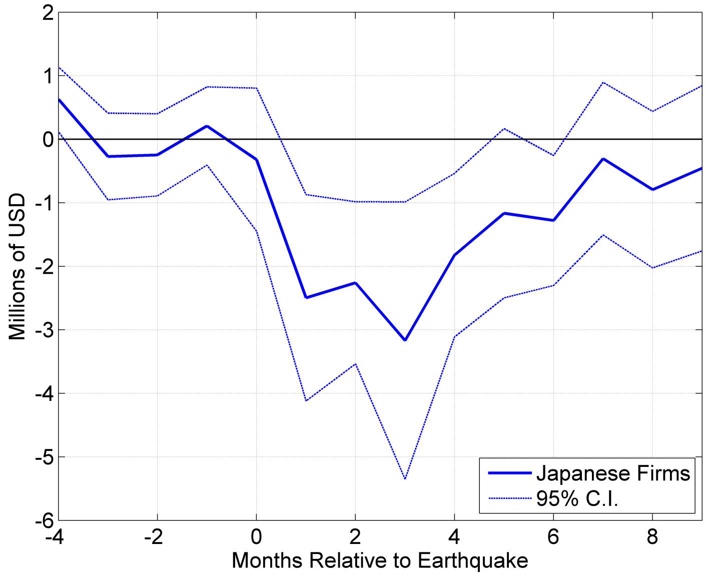
Figure 3.5: Density of Firm-Level Exposure to Japanese Imported Inputs: By Firm Type



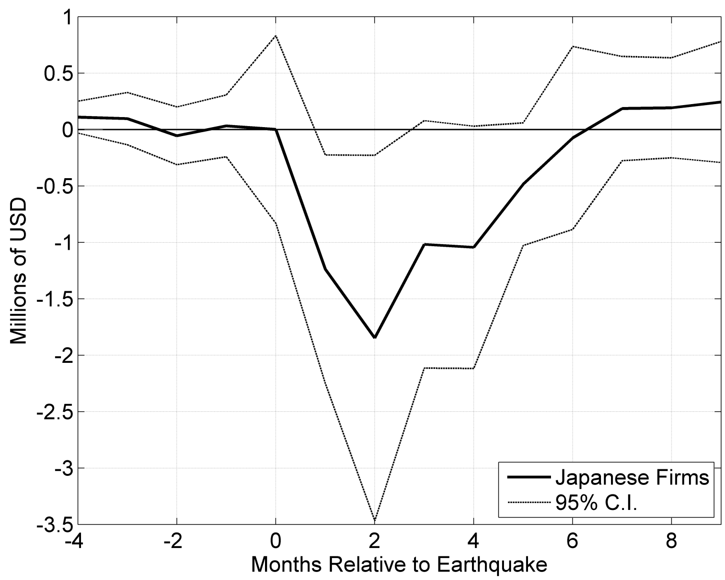
Source: LFTTD-DCA-UBP as explained in text. The estimates correspond to year 2010. This figure displays density estimates of the log exposure measure to Japanese imported inputs, separately for Japanese affiliates and non-Japanese multinational firms. The measure is defined as the ratio of Japanese imported inputs to total imported inputs plus U.S. salaries and wages. Estimates at either tail are suppressed for confidentiality purposes.

Figure 3.6: Dynamic Treatment Effects: Japanese Firms

A. Relative Intermediate Input Imports of Japanese Firms

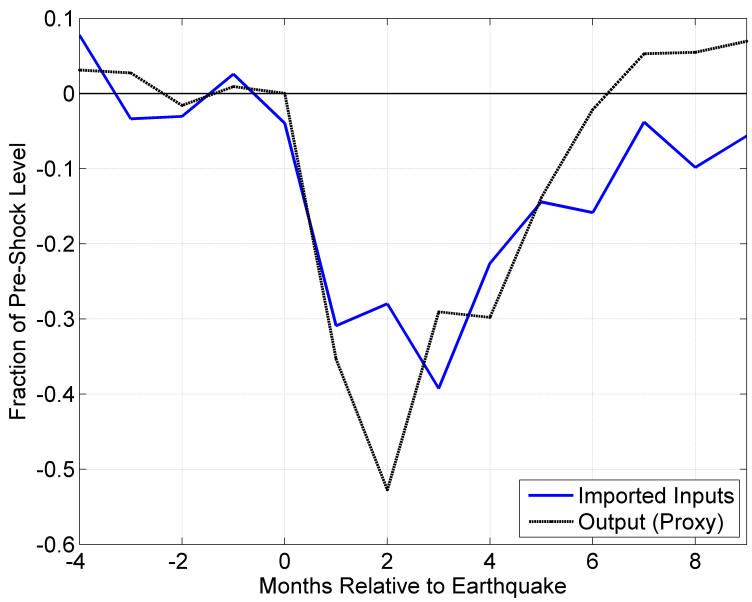


B. Relative North American Exports of Japanese Firms



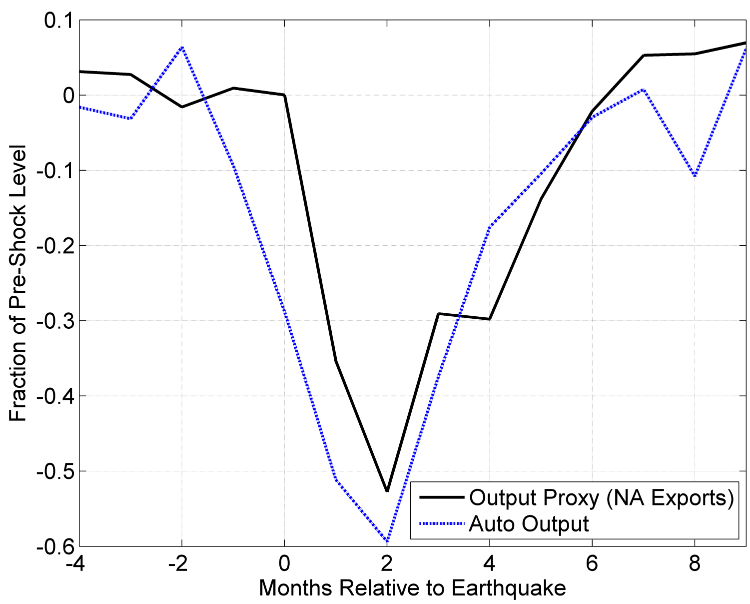
Source: LFTTD-DCA-UBP as explained in text.
These figures report the intermediate imports and North American exports of the U.S. affiliates of Japanese firms relative to a control group of other multinational firms. The values are coefficient estimates taken from an interaction of a Japanese-firm dummy with a monthly dummy – additional baseline monthly dummies remove seasonal effects. See equation 3.5 in the text. Standard errors are clustered at the firm level.

Figure 3.7: Relative Imported Inputs and Output (Proxy) of Japanese Firms: Fraction of Pre-Shock Level



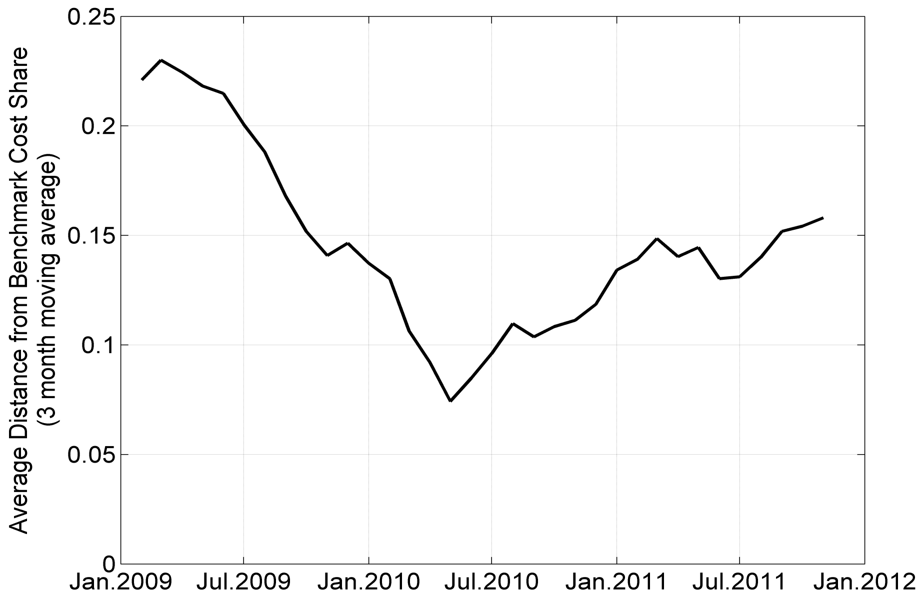
Source: LFTTD-DCA-UBP as explained in text.
This figure reports the intermediate imports and output proxy (North American exports) of the U.S. affiliates of Japanese firms relative to a control group of other multinational firms. The values are percent changes from the pre-shock level of each series, defined as the average of the months December 2010, January 2011, and February 2011.

Figure 3.8: Assessing the Output Proxy Using Monthly Automotive Production



Source: Ward's Automotive Database
This figure reports the production levels of Japanese auto plants relative to a control group of non-Japanese auto plants. The values are percent changes from a pre-shock level, defined as the average of the months December 2010, January 2011, and February 2011. See equation C.14 in the text. For purposes of comparison, we also include the equivalent measure corresponding to total manufacturing of Japanese affiliates using the output proxy from Census data (from Figure 3.7). The Japanese automakers are Honda, Mazda, Mitsubishi, Nissan, Toyota, and Subaru. For the sake of clarity, we suppress the standard errors for the automotive series, though there are 4 months with below zero production based on a 95 percent confidence interval. See Appendix C.3.7 for more details.

Figure 3.9: Japanese Products: Average Distance from Benchmark Cost Shares: JPN Multinationals



Source: LFTTD-DCA-UBP as explain in the text
Underlying this figure is the calculation of the average total (absolute) deviations from a benchmark measure of a firm's cost shares across input products from Japan. See equation 3.17 in the text. The figure reports the mean across the Japanese multinationals used in the section 3.4.

APPENDICES

APPENDIX A

Chapter 1 Appendices

A.1 Model appendix

A.1.1 Summary of equations

In this part of the Appendix I summarize the equations of the most general model with adjustment costs and time-varying discount factor. $q_{X,t}$ and $q_{Z,t}$ denote the shadow values of one unit of capital in the two sectors. To obtain the model in Section 1.2 set $\beta_t = \beta$ and $\zeta_K = 0$.

The household's behavior is summarized the following equations:

$$\lambda_t = (1 + i_t)\, \mathbb{E}_t \left[\beta_{t+1}\lambda_{t+1}\right] \tag{A.1}$$

$$\frac{\partial v\left(N_{Z,t}, N_{X,t}\right)}{\partial N_{j,t}} = \lambda_t W_{j,t},\ j \in \{X, Z\} \tag{A.2}$$

$$\frac{\partial u\left(C_t, D_{H,t}\right)}{\partial C_t} = \lambda_t P_{Z,t} \tag{A.3}$$

$$q_{j,t} = \lambda_t P_{Z,t}\left(1 + \zeta_K\left(\frac{I_{j,t}}{K_{j,t-1}} - \delta_K\right)\right),\ j \in \{X, Z\} \tag{A.4}$$

$$q_{j,t}=\mathbb{E}_t\left[\beta_{t+1}\lambda_{t+1}R_{j,t+1}+\beta_{t+1}\lambda_{t+1}P_{j,t+1}\frac{\zeta_K}{2}\left(\left(\frac{I_{j,t+1}}{K_{j,t}}\right)^2-\left(\delta_K\right)^2\right)+\beta_{t+1}\left(1-\delta_K\right)q_{j,t+1}\right],\ j\in\{X,Z\} \tag{A.5}$$

$$\gamma_t=\frac{\partial u\left(C_t,D_{H,t}\right)}{\partial D_{H,t}}+\left(1-\delta_D\right)\mathbb{E}_t\left[\beta_{t+1}\gamma_{t+1}\right]. \tag{A.6}$$

The accumulation equations (1.2) and equation (1.15) in the text continue to hold.

Firms' reset their prices according to

$$p^*_{j,t}=\frac{\varepsilon}{\varepsilon-1}\frac{\mathbb{E}_t\sum_{k=0}^{\infty}\left(\theta_j\right)^k\lambda_{t+k}\prod_{s=1}^{k}\beta_{t+s}\left(P_{j,t+k}\right)^{\varepsilon}j_{t+k}MC_{j,t+k}}{\mathbb{E}_t\sum_{k=0}^{\infty}\left(\theta_j\right)^k\lambda_{t+k}\prod_{s=1}^{k}\beta_{t+s}\left(P_{j,t+k}\right)^{\varepsilon}j_{t+k}},\ j\in\{X,Z\} \tag{A.7}$$

and prices in both sectors evolve according to

$$P_{j,t}=\left(\theta_j\left(P_{j,t-1}\right)^{1-\varepsilon}+\left(1-\theta_j\right)\left(p^*_{j,t}\right)^{1-\varepsilon}\right)^{\frac{1}{1-\varepsilon}},\ j\in\{X,Z\}. \tag{A.8}$$

Nominal marginal costs in the two sectors are

$$MC_{X,t}\left(s\right)=\left(\frac{P_{Z,t}}{1-\chi}\right)^{1-\chi}\left(\frac{R_{X,t}}{\alpha\chi}\right)^{\chi\alpha}\left(\frac{W_{X,t}}{\left(1-\alpha\right)\chi}\right)^{\chi(1-\alpha)}, \tag{A.9}$$

$$MC_{Z,t}\left(s\right)=\left(\frac{R_{Z,t}}{\alpha}\right)^{\alpha}\left(\frac{W_{Z,t}}{1-\alpha}\right)^{1-\alpha}. \tag{A.10}$$

The conditional factor demand functions are

$$M_{X,t}=\left(\frac{P_{Z,t}}{1-\chi}\right)^{-\chi}\left(\frac{R_{X,t}}{\alpha\chi}\right)^{\chi\alpha}\left(\frac{W_{X,t}}{\left(1-\alpha\right)\chi}\right)^{\chi(1-\alpha)}X_t, \tag{A.11}$$

$$K_{X,t-1}=\left(\frac{P_{Z,t}}{1-\chi}\right)^{1-\chi}\left(\frac{R_{X,t}}{\alpha\chi}\right)^{\chi\alpha-1}\left(\frac{W_{X,t}}{\left(1-\alpha\right)\chi}\right)^{\chi(1-\alpha)}X_t, \tag{A.12}$$

$$N_{X,t}=\left(\frac{P_{Z,t}}{1-\chi}\right)^{1-\chi}\left(\frac{R_{X,t}}{\alpha\chi}\right)^{\chi\alpha}\left(\frac{W_{X,t}}{\left(1-\alpha\right)\chi}\right)^{\chi(1-\alpha)-1}X_t, \tag{A.13}$$

$$K_{Z,t-1} = \left(\frac{R_{Z,t}}{\alpha}\right)^{\alpha-1} \left(\frac{W_{Z,t}}{1-\alpha}\right)^{1-\alpha} Z_t, \tag{A.14}$$

$$N_{Z,t} = \left(\frac{R_{Z,t}}{\alpha}\right)^{\alpha} \left(\frac{W_{Z,t}}{1-\alpha}\right)^{-\alpha} Z_t. \tag{A.15}$$

The market clearing condition in the Z sector with capital adjustment costs is

$$Z_t = C_t + I_{X,t} + I_{Z,t} + M_{X,t} + K_{X,t-1}\frac{\zeta_K}{2}\left(\frac{I_{X,t}}{K_{X,t-1}} - \delta_K\right)^2 + K_{Z,t-1}\frac{\zeta_K}{2}\left(\frac{I_{Z,t}}{K_{Z,t-1}} - \delta_K\right)^2. \tag{A.16}$$

Market clearing for the X sector, accounting, government spending, and the monetary policy rule are as described in the text (equations 1.7 to 1.13).

A.1.2 Proofs of approximation results

In what follows the notation $\tilde{X}_t = \frac{X_t - X}{X}$ denotes the percentage deviation of variable X_t from its steady state value. I prove the results for the case in which the fiscal policy shock has no persistence, $\varrho_X = 0$, although they can be generalized to cases with mild persistence. I assume that prior to the shock the economy is in the steady state. Notice that the references in Appendix A.1.1 correspond to the baseline model in the text when there are no adjustment costs, $\zeta_K = 0$, and the discount factor β is constant.

A.1.2.1 Spending on durable goods

Approximation result 1. *Suppose δ_K and δ_D are arbitrarily close to zero and β is arbitrarily close to* 1. *Then, for a short-lived increase in spending, it is approximately true that (1) $\Delta X_{H,t} \approx -\Delta X_{G,t}$, (2) the price $P_{X,t}$ remains unchanged, (3) the sectoral multipliers for gross output and value added are zero, $\frac{\mathrm{d}GO_{X,t}}{\mathrm{d}G_{X,t}} \approx \frac{\mathrm{d}VA_{X,t}}{\mathrm{d}G_{X,t}} \approx 0$, and (4) the aggregate multiplier is zero, $\frac{\mathrm{d}Y_t}{\mathrm{d}G_{X,t}} \approx 0$.*

Proof I first show that these assumptions imply that the stocks $K_{X,t}, K_{Z,t}$ and $D_{H,t}$

are approximately constant. The linear approximation of the accumulation equation for durables (equation 1.2) is

$$\tilde{D}_{H,t} = \delta_D \tilde{X}_{H,t} + (1 - \delta_D) \tilde{D}_{H,t-1}$$

It then follows from the assumptions $\delta_D \to 0$ and $\tilde{D}_{H,t-1} = 0$ that $\tilde{D}_{H,t} \approx 0$. Similarly, $\tilde{K}_{X,t} \approx \tilde{K}_{Z,t} \approx 0$.

The respective shadow values γ_t, $q_{X,t}$, $q_{Z,t}$ are also approximately constant under these assumptions. This can be seen using the linear approximation of equation (A.6),

$$\tilde{\gamma}_t = -\frac{(1 - (1 - \delta_D) \beta)}{\sigma_D} \tilde{D}_{H,t} + \frac{(1 - (1 - \delta_D) \beta)}{\sigma_{DC}} \tilde{C}_t + (1 - \delta_D) \beta \mathbb{E}_t [\tilde{\gamma}_{t+1}]$$

Here, σ_D and σ_{DC} are constants. Clearly, under the assumptions $\delta_D \to 0$ and $\beta \to 1$ it follows that $\tilde{\gamma}_t \approx \mathbb{E}_t [\tilde{\gamma}_{t+1}]$. Since this variable is stationary and returns to its steady state value eventually it must be that $\tilde{\gamma}_t \approx 0$. Similarly, $q_{X,t} \approx q_{Z,t} \approx 0$.

I next guess that $\tilde{P}_{X,t} = 0$. It then follows from equations (1.15) and (A.4) that $\tilde{\lambda}_t = 0$ and $\tilde{P}_{Z,t} = 0$. This guess, together with $\tilde{D}_{H,t} \approx \tilde{K}_{X,t} \approx \tilde{K}_{Z,t} \approx 0$ and the assumption that $\varrho_X = 0$ implies that all state variables from period $t + 1$ onwards are zero. Hence all other variables from period $t + 1$ onwards are zero.

With these results in hand it is easy to show that

$$\Delta X_{H,t} = -\Delta X_{G,t},$$

$\Delta I_{X,t} + \Delta I_{Z,t} = 0$ and that no other variable at time t responds to the fiscal policy shock. Notice that since neither of the prices adjust nor the quantities X_t, Z_t or Y_t change, the monetary policy rule (1.13) implies that the nominal interest rate remains unchanged. The claims on the multipliers now follow immediately.

A.1.2.2 Spending on nondurable goods

Approximation result 2. *Suppose $\delta_D = 1$, δ_K is arbitrarily close to zero and β is arbitrarily close to 1. Suppose further that $\theta_X = \theta_Z = 0$, that is, prices are fully flexible, and that $u_{CD} = v_{XZ} = 0$. Lastly, assume that the government subsidizes monopolistic firms such that $\frac{P_Z M_X}{P_X X} = 1 - \chi$. Next, define*

$$\sigma_D = -\frac{u_D(C, D_H)}{D_H u_{DD}(C, D_H)} \text{ and } \eta_X = \frac{v_X(N_Z, N_X)}{N_X v_{XX}(N_Z, N_X)}.$$

Then a short-lived increase in spending yields a gross output sectoral multiplier equal to

$$\frac{\mathrm{d}GO_{X,t}}{\mathrm{d}G_{X,t}} \approx \frac{\left(1+\eta_X^{-1}\right)\chi^{-1} - \left(\alpha+\eta_X^{-1}\right)}{\sigma_D\left(\alpha+\eta_X^{-1}\right)\frac{X_H}{X} + \left(1+\eta_X^{-1}\right)\chi^{-1} - \left(\alpha+\eta_X^{-1}\right)}$$

and approximately equal sectoral value added and aggregate multipliers

$$\frac{\mathrm{d}VA_{X,t}}{\mathrm{d}G_{X,t}} \approx \frac{\mathrm{d}Y_t}{\mathrm{d}G_{X,t}} \approx \frac{1-\alpha}{\sigma_D\left(\alpha+\eta_X^{-1}\right)\frac{X_H}{X} + \left(1+\eta_X^{-1}\right)\chi^{-1} - \left(\alpha+\eta_X^{-1}\right)}.$$

The relative price $P_{X,t}/P_{Z,t}$ rises in response to greater spending.

Proof Similar to the proof in A.1.2.1, the assumptions that $\beta \to 1$ and $\delta_K \to 0$ implies that $\tilde{K}_{X,t} \approx \tilde{K}_{Z,t} \approx \tilde{q}_{Z,t} \approx \tilde{q}_{X,t} \approx 0$. It then immediately follows from (A.4) that $\tilde{\lambda}_t = -\tilde{P}_{Z,t}$. Additionally, equation (1.15) implies that $\tilde{\gamma}_t = \tilde{P}_{X,t} - \tilde{P}_{Z,t}$, and (A.3) implies that $\tilde{C}_t = 0$. Using these relationships in the labor supply functions (A.2) yields

$$\frac{1}{\eta_Z}\tilde{N}_{Z,t} = \tilde{W}_{Z,t} - \tilde{P}_{Z,t} \text{ and } \frac{1}{\eta_X}\tilde{N}_{X,t} = \tilde{W}_{X,t} - \tilde{P}_{Z,t}.$$

Additionally, the household's demand for good X (equation A.6) becomes

$$\tilde{P}_{X,t}-\tilde{P}_{Z,t}=-\frac{1}{\sigma_D}\tilde{X}_{H,t} \tag{A.17}$$

With these results in hand, it is straightforward but tedious to show that $\tilde{Z}_t=0$ and that

$$\tilde{X}_t=\frac{\left(1+\eta_X^{-1}\right)\chi^{-1}-\left(\alpha+\eta_X^{-1}\right)}{\left(1-\chi\alpha+(1-\chi)\,\eta_X^{-1}\right)\chi^{-1}+\left(\alpha+\eta_X^{-1}\right)\sigma_D\frac{X_H}{X}}\frac{X_G}{X}\tilde{X}_{G,t}$$

and

$$\tilde{M}_{X,t}=\frac{\left(1+\eta_X^{-1}\right)\chi^{-1}}{\left(1-\chi\alpha+(1-\chi)\,\eta_X^{-1}\right)\chi^{-1}+\left(\alpha+\eta_X^{-1}\right)\sigma_D\frac{X_H}{X}}\frac{X_G}{X}\tilde{X}_{G,t}$$

Recall the assumption that there is a subsidy that implies that $\frac{P_Z M_X}{P_X X}=1-\chi$. An ad-valorem subsidy of $\frac{1}{\varepsilon}$ on purchases of the intermediate is one way to ensure this relationship. Using the definitions of gross output, value added, and GDP, and the fact that $\frac{P_Z M_X}{P_X X}=1-\chi$ the results for the multipliers are now easily shown. Notice finally that $\tilde{P}_{X,t}-\tilde{P}_{Z,t}$ rises as $\tilde{X}_{H,t}$ falls (equation A.17).

A.1.3 Robustness of fiscal multipliers to alternative calibrations

In this part of the appendix I check the robustness of the fiscal multipliers to alternative calibrations. The first three rows in Table A.1 illustrate the role of capital adjustment costs. Whereas neither gross output nor value added sectoral multipliers change substantially with adjustment costs, aggregate multipliers do. Higher adjustment costs imply less crowding out of investment through purchases of intermediates. When adjustment costs are high, aggregate multipliers are quite close to the gross output multipliers.

I next relax the assumption that labor is immobile across sectors. Specifically, I assume that

$$v\left(N_{X,t},N_{Z,t}\right)=\left(1+\frac{1}{\eta}\right)^{-1}\left[\left(\phi\left(N_{Z,t}\right)^{\frac{\eta+\mu}{\eta}}+(1-\phi)\left(N_{X,t}\right)^{\frac{\eta+\mu}{\eta}}\right)^{\frac{\eta}{\eta+\mu}}\right]^{\left(1+\frac{1}{\eta}\right)},$$

Here, η is the Frisch labor supply elasticity and μ parameterizes labor mobility across sectors. If $\mu = 0$, labor is perfectly mobile across sectors. If $\mu = 1$, labor is completely immobile. Intermediate values imply partial labor mobility.[1]

Since greater labor mobility raises the elasticity of both sectors' supply curves, the sectoral multipliers increase slightly. Of course, more workers in one sector imply fewer in the other so the effect on the aggregate multiplier is small. In order to mimic severe slack in labor markets in recessions, I also compute multipliers for a calibration with greater labor supply elasticity ($\eta = 2$). This parameterization also implies more elastic sectoral supply curves but now additional hires in one sector do not draw labor away from the other. As a result all multipliers increase relative to the baseline.

An additional determinant of the short-run elasticity of supply curves are sticky prices. If prices are sticky, as in the baseline calibration, a fraction of firms must serve increased demand at fixed prices. When I assume that prices are flexible, the case $\theta_X = \theta_Z = 0$, it is therefore unsurprising that multipliers fall. Naturally, this decline in multipliers relative to the baseline calibration is larger in the more price sensitive durable goods sectors.

Next, I consider an alternative Taylor rule in which the monetary authority responds stronger to both inflation and deviations of output from trend ($\phi_\pi = 1.5$, $\phi_Y = 0.5/4$). As the open economy relative multiplier in *Nakamura and Steinsson* (2014), sectoral multipliers are less sensitive to alternative monetary policy rules than aggregate multipliers. Of course, aggregate multipliers fall when the monetary authority "leans against the wind" with increasing strength.

I next assume that the persistence of the government spending shock is higher and takes a value of 0.90. This value corresponds to the persistence estimated in section 1.3.2.3 ($0.9^{4\cdot 7} \approx 0.05$). Longer lived shocks raise the multipliers for durable goods and reduces those for nondurables. Yet, the aggregate nondurables multiplier with a value of 0.70 remains more

[1]This specification is taken from *Barsky et al.* (2003).

than twice the size of the durables multiplier which takes a value of 0.32. For the empirical analysis in this paper, it is important that the industry-level multipliers also remain very different.

Finally, I explore the impact on fiscal multipliers when D_H and C are complements or substitutes (the last two lines of Table A.1). The more complementary D_H and C are, the larger are the durable goods multipliers. However, the difference between nondurables and durable's multipliers remains large.

Across all calibrations, a good rule of thumb is that the aggregate multipliers lie between sectoral value added and gross output multipliers. In the simple model presented here, this rule suggests that the aggregate multiplier is between 0.1 and 0.35 for durable goods and around 0.75 for nondurables goods.

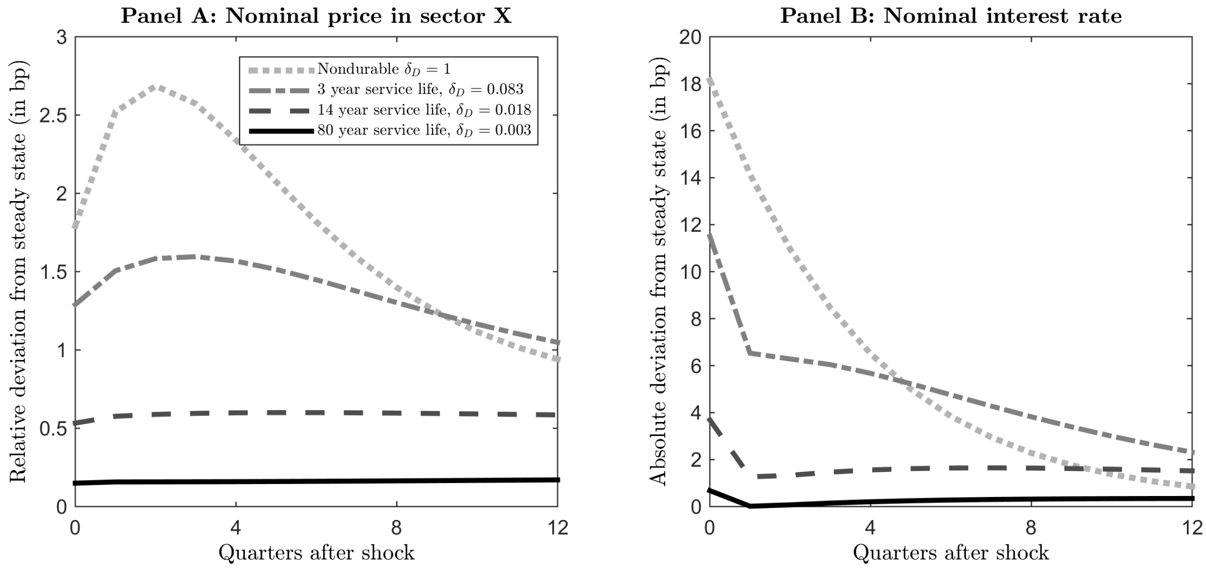
Table A.1: Multipliers for alternative calibrations

Small sector	15 year service life, $\delta_D = 0.017$			Nondurable $\delta_D = 1$		
	Sectoral multiplier			Sectoral multiplier		
Calibration	Gross output	Value added	Aggregate multiplier	Gross output	Value added	Aggregate multiplier
Baseline (baseline calibration)	0.22	0.08	0.21	0.88	0.30	0.76
High adjustment costs $\zeta_K = 100$	0.22	0.08	0.22	0.88	0.30	0.83
Low adjustment costs $\zeta_K = 1$	0.20	0.08	0.18	0.88	0.32	0.43
High labor mobility $\mu = 0.5$	0.25	0.10	0.23	0.90	0.35	0.75
High labor supply elasticity $\eta = 2$	0.25	0.10	0.26	0.90	0.35	0.85
Flexible prices $\theta_X = \theta_Z = 0$	0.10	0.04	0.13	0.75	0.26	0.60
More aggressive Taylor rule $\phi_\pi = 1.5,\ \phi_Y = 0.5/4$	0.22	0.08	0.20	0.87	0.30	0.63
Greater shock persistence $\varrho_X = 0.90$	0.34	0.12	0.32	0.86	0.30	0.70
D_H and C complements $\rho = 0.5$	0.33	0.11	0.29	0.93	0.32	0.76
D_H and C substitutes $\rho = 2$	0.13	0.05	0.15	0.79	0.27	0.75

Notes: The table reports the multipliers at a time horizon of one year after the shock as generated by the model.

A.1.4 Additional impulse response functions

Figure A.1: Impulse response functions for a government spending shock



Notes: The figure plots impulse response functions for various calibrations. The impulse is a 100 basis point increase of government spending in sector X relative to its steady state value.

A.2 Appendix for aggregate empirical analysis

A.2.1 Data used in aggregate analysis

The data are annual and come from the Bureau of Economic Analysis. GDP, the consumption series and investment in current dollars are taken from Table 1.1.5 and the corresponding price indexes are taken from Table 1.1.4. For $G_{X,t}$ I use national defense gross investment, and for $G_{C,t}$ I use national defense consumption expenditures. These series in current dollars are obtained from Table 3.9.5 and the price indexes from Table 3.9.4. Note that consumption expenditures include consumption of fixed capital. Additionally, roughly 5 to 10 percent of defense consumption expenditures are purchases of durable goods which cannot be separated out until 1972 because the Bureau of Economic Analysis begins to report a detailed breakdown of military spending only in 1972, Table 3.11.5. Hence, a small component of durable goods spending remains in the series $G_{C,t}$. Valerie Ramey's news series can be downloaded from her website (`http://econweb.ucsd.edu/~vramey/research.html#data`). I obtain the trend of GDP Y_t^T using an HP-filter with a smoothing parameter of 6.25 as recommended by *Ravn and Uhlig* (2002).

A.2.2 Additional evidence from national accounts data

From 1972 onwards, the BEA provides a more detailed breakdown of military spending in Table 3.11.5. This table contains time series of defense spending separately for services, nondurable goods, durable goods, equipment, and structures, among others.[2] I can use this data to test whether spending on these five categories leads to crowding out in the corresponding spending category for the private sector. To do so, I group services and nondurable goods together into a nondurables group, denoted C, and durable goods, equipment and structures into a durables group X. The data are now quarterly.

Let $Y_{P,i,t}$ denote private sector spending on category i. I estimate the specification

$$\frac{Y_{P,i,t} - Y_{P,i,t-1}}{Y^T_{P,i,t-4}} = \sum_{s=1}^{4} \gamma_{X,s} \frac{G_{X,i,t-s+1} - G_{X,i,t-s}}{Y^T_{P,i,t-4}} + \sum_{s=1}^{4} \gamma_{C,s} \frac{G_{C,i,t-s+1} - G_{C,i,t-s}}{Y^T_{P,i,t-4}} + \delta_t + \zeta_i + \varepsilon_{i,t}. \tag{A.18}$$

This empirical model is closely related to equation (1.19), but adjusted for quarterly data. Instead of Ramey's news variable, I now include a time fixed effect, δ_t, which also soaks up the common component of other shocks such as monetary and tax policy. ζ_i is a fixed effect for each spending category. $G_{X,i,t}$ ($G_{C,i,t}$) denotes defense spending on category i interacted with a dummy variable that takes the value one if and only if industry i belongs to the durables group (the nondurables group). The multipliers are simply $\sum_{s=1}^{4} \gamma_{j,s}$, $j = C, X$.

[2]The deflators are in Table 3.11.4.

Table A.2: Estimates from disaggregated national accounts data

Specification	(1)	(2)	(3)
Durable goods multiplier	-1.12 (0.46)	-1.04 (0.47)	-1.17 (0.47)
Nondurable goods multiplier	3.99 (2.88)	4.34 (2.86)	4.32 (2.94)
Controls in addition to those in equation (A.18)	none	interest rate interacted with category dummy	Ramey news variable interacted with category dummy
Observations	845	845	845

Note: Standard errors in parentheses.

Table A.2 shows the results. In the baseline specification, column (1), the point estimate is -1.12, suggesting that one dollar of spending in a durables category (durable goods, equipment, and structures) crowds out 1.12 dollars of private sector spending. The estimate is significantly different from zero at the 5 percent level. In contrast, the estimate for nondurable goods is positive and insignificantly different from zero. Unfortunately, the standard errors are fairly large.

In columns (2) and (3) of the table I report two estimates from specifications that include additional control variables. Column (2) reports the estimates when I include the real interest rate interacted with a category dummy. This control variable would allow monetary policy to affect the spending categories differently. The results in column (3) are obtained when I additionally control for Ramey's news variable interacted with category dummies. As can be seen from the table, the results are not sensitive to the inclusion of these additional controls. Although I do not report the estimates here, the findings are also robust to the inclusion of additional lags in specification (A.18) so that the time horizon over which the multipliers are computed is greater than four quarters.

A.3 Appendix for industry-level empirical analysis

A.3.1 Industry-level data

A.3.1.1 Original data sources for sectoral analysis

The NBER-CES manufacturing database can be downloaded from `http://www.nber.org/nberces/`. This page also provides summary statistics of the dataset and a documentation. As noted in the text, a detailed description of the database is provided by *Bartelsman and Gray* (1996) and *Becker et al.* (2013).

The military prime contract files were downloaded from the U.S. National Archives, `https://research.archives.gov/`. After entering the website, search for the key words "military prime contracts" and "defense contract action data system". There is a separate set of files, i.e. the dataset and documentation, for each fiscal year from 1966 to 2003. There is also a separate file for the 3-month period between the old and the new fiscal year in 1975.

The remaining data on military spending were obtained from `www.usaspending.gov`. More precisely, the data can be downloaded from `https://www.usaspending.gov/DownloadCenter/Pages/DataDownload.aspx` for the years from 2000 onwards. When downloading the data be sure to select "prime award" as type of data and "contracts" as spending type. The agency is the Department of Defense.

A.3.1.2 Product Service Codes and Federal Supply Codes

U.S. military procurement is categorized according to Product Service Codes (PSCs) and Federal Supply Codes (FSCs). Several examples are given in table A.3. Individual categories can be looked up here: `http://support.outreachsystems.com/resources/tables/pscs/`. A summary of all FSC and PSC codes as used after 1979 is available here: `https://www.fpds.gov/downloads/psc_data_Oct012011.xls`. Both websites were last accessed on 06/16/2015.

Table A.3: Selected examples of PSC and FSC codes

PSC/FSC	Description
1010	Guns, over 30mm up to 75mm
1560	Airframe structural components
1615	Helicopter rotor blades, drive mechanisms and components
3820	Mining, rock drilling, earth boring, and related equipment
AR32	R&D-Space: Flight (applied research/exploratory development)
M1GC	Operation of fuel storage buildings
R702	Data collection services

A.3.1.3 Concordance between FSC/PSC and SIC codes

The military prime contract files contain both FSC/PSC codes and 4-digit SIC codes for the period from 1989 to 2000. I use these 12 years to construct a concordance and then apply the concordance to spending on FSC/PSC categories which are available over the entire sample from 1979 to 2009.

The concordance is a matrix that describes for each FSC and PSC code what fraction of a dollar spend on the FSC/PSC code is purchased from each SIC industry. For instance, if one dollar is spent on the FSC code 1010 (Guns, over 30mm up to 75mm) about 45 cents are purchased from SIC industry 3484 (Small Arms). The next most important SIC industry is 3489 (Ordnance and Accessories, NEC) with 25 cents. Table A.4 provides a summary of all SIC industries that receive more than 1 cent when one dollar is spent on FSC code 1010. A second example is given in Table A.5 for the FSC code 1560 (Airframe Structural Components).

Table A.4: Spending shares for FSC code 1010 (Guns, over 30mm up to 75mm)

SIC code	Description	Spending share
3484	Small Arms	0.448
3489	Ordnance and Accessories, NEC	0.254
3499	Fabricated Metal Products, NEC	0.166
8711	Engineering Services	0.086
3728	Aircraft Parts and Auxiliary Equipment, NEC	0.021

Table A.5: Spending shares for FSC code 1560 (Airframe Structural Components)

SIC code	Description	Spending share
3728	Aircraft Parts and Auxiliary Equipment, NEC	0.815
3721	Aircraft	0.069
8711	Engineering Services	0.033
3724	Aircraft Engines and Engine Parts	0.023
8731	Commercial Physical and Biological Research	0.022

A.3.2 Sample description

Table A.6: Industries and basic summary statistics

SIC code	Durable	Avg. def. spending share (in %)	Description
2032	no	1.8	Canned Specialties
2086	no	2.0	Bottled and Canned Soft Drinks and Carbonated Water
2097	no	1.0	Manufactured Ice
2099	no	1.2	Food Preparations, NEC
2231	no	5.7	Broadwoven Fabric Mills, Wool
2254	no	3.1	Knit Underwear and Nightwear Mills
2298	no	2.2	Cordage and Twine
2299	no	2.7	Textile Goods, NEC
2311	no	9.0	Mens and Boys Suits, Coats, and Overcoats
2321	no	1.9	Mens and Boys Shirts, Except Work Shirts
2322	no	10.9	Mens and Boys Underwear and Nightwear
2325	no	7.3	Mens and Boys Separate Trousers and Slacks
2326	no	3.8	Mens and Boys Work Clothing
2329	no	6.8	Mens and Boys Clothing, NEC
2353	no	6.1	Hats, Caps, and Millinery
2371	no	2.8	Fur Goods
2381	no	17.6	Dress and Work Gloves, Except Knit and All-Leather
2385	no	32.2	Waterproof Outerwear
2387	no	1.5	Apparel Belts
2389	no	2.6	Apparel and Accessories, NEC
2393	no	5.9	Textile Bags
2394	no	8.7	Canvas and Related Products
2399	no	8.7	Fabricated Textile Products, NEC
2519	yes	1.1	Household Furniture, NEC

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Table A.6 – *Continued from previous page*

SIC code	Durable	Avg. def. spending share (in %)	Description
2521	yes	2.4	Wood Office Furniture
2522	yes	2.0	Office Furniture, Except Wood
2599	yes	1.5	Furniture and Fixtures, NEC
2741	no	2.1	Miscellaneous Publishing
2813	no	2.4	Industrial Gases
2836	no	1.4	Biological Products, Except Diagnostic Substances
2892	no	22.2	Explosives
2911	no	1.6	Petroleum Refining
2992	no	1.7	Lubricating Oils and Greases
3021	no	2.5	Rubber and Plastics Footwear
3053	no	1.3	Gaskets, Packing, and Sealing Devices
3069	no	1.4	Fabricated Rubber Products, NEC
3143	no	6.9	Mens Footwear, Except Athletic
3149	no	5.5	Footwear, Except Rubber, NEC
3151	no	12.1	Leather Gloves and Mittens
3261	yes	2.3	Vitreous China Plumbing Fixtures and China and Earthenware Fittings and Bathroom Accessories
3295	yes	5.3	Minerals and Earths, Ground or Otherwise Treated
3299	yes	1.1	Nonmetallic Mineral Products, NEC
3315	yes	1.1	Steel Wiredrawing and Steel Nails and Spikes
3399	yes	1.4	Primary Metal Products, NEC
3412	yes	1.6	Metal Shipping Barrels, Drums, Kegs, and Pails
3429	yes	6.2	Hardware, NEC
3443	yes	6.1	Fabricated Plate Work
3448	yes	1.9	Prefabricated Metal Buildings and Components
3452	yes	1.6	Bolts, Nuts, Screws, Rivets, and Washers

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Table A.6 – *Continued from previous page*

SIC code	Durable	Avg. def. spending share (in %)	Description
3484	yes	16.3	Small Arms
3494	yes	6.4	Valves and Pipe Fittings, NEC
3499	yes	4.3	Fabricated Metal Products, NEC
3511	yes	8.6	Steam, Gas, and Hydraulic Turbines, and Turbine Generator Set Units
3519	yes	1.6	Internal Combustion Engines, NEC
3536	yes	3.8	Overhead Traveling Cranes, Hoists, and Monorail Systems
3537	yes	5.7	Industrial Trucks, Tractors, Trailers, and Stackers
3541	yes	1.6	Machine Tools, Metal Cutting Types
3542	yes	1.2	Machine Tools, Metal Forming Type
3559	yes	4.8	Special Industry Machinery, NEC
3561	yes	1.6	Pumps and Pumping Equipment
3562	yes	1.7	Ball and Roller Bearings
3564	yes	1.2	Industrial and Commercial Fans and Blowers and Air Purification Equipment
3566	yes	6.5	Speed Changers, Industrial High-Speed Drives, and Gears
3568	yes	2.9	Mechanical Power Transmission Equipment, NEC
3569	yes	2.7	General Industrial Machinery and Equipment, NEC
3571	yes	4.2	Electronic Computers
3572	yes	1.3	Computer Storage Devices
3575	yes	11.2	Computer Terminals
3577	yes	3.1	Computer Peripheral Equipment, NEC
3578	yes	1.1	Calculating and Accounting Machinery, Except Electronic Computers
3579	yes	2.1	Office Machines, NEC
3582	yes	1.5	Commercial Laundry, Drycleaning, and Pressing Machines
3586	yes	1.3	Measuring and Dispensing Pumps

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Table A.6 – *Continued from previous page*

SIC code	Durable	Avg. def. spending share (in %)	Description
3589	yes	1.2	Service Industry Machinery, NEC
3612	yes	3.7	Power, Distribution, and Specialty Transformers
3613	yes	1.0	Switchgear and Switchboard Apparatus
3621	yes	3.5	Motors and Generators
3629	yes	1.8	Electrical Industrial Apparatus, NEC
3643	yes	1.6	Current-Carrying Wiring Devices
3644	yes	1.7	Noncurrent-Carrying Wiring Devices
3647	yes	1.7	Vehicular Lighting Equipment
3648	yes	1.3	Lighting Equipment, NEC
3661	yes	2.0	Telephone and Telegraph Apparatus
3663	yes	11.6	Radio and Television Broadcasting and Communications Equipment
3671	yes	23.1	Electron Tubes
3672	yes	2.7	Printed Circuit Boards
3676	yes	3.0	Electronic Resistors
3677	yes	1.5	Electronic Coils, Transformers, and Other Inductors
3679	yes	6.6	Electronic Components, NEC
3691	yes	1.5	Storage Batteries
3692	yes	4.4	Primary Batteries, Dry and Wet
3695	yes	2.8	Magnetic and Optical Recording Media
3699	yes	10.8	Electrical Machinery, Equipment, and Supplies, NEC
3715	yes	4.3	Truck Trailers
3721	yes	29.4	Aircraft
3724	yes	27.6	Aircraft Engines and Engine Parts
3732	yes	1.2	Boat Building and Repairing
3764	yes	12.5	Guided Missile and Space Vehicle Propulsion Units and Propulsion Unit Parts

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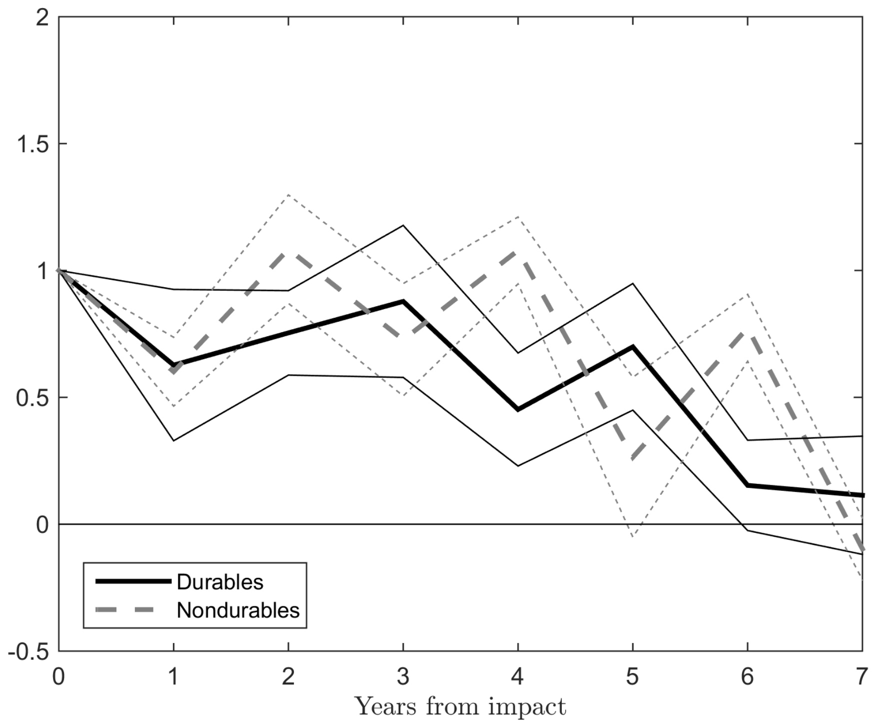
Table A.6 – *Continued from previous page*

SIC code	Durable	Avg. def. spending share (in %)	Description
3799	yes	1.5	Transportation Equipment, NEC
3812	yes	15.1	Search, Detection, Navigation, Guidance, Aeronautical, and Nautical Systems and Instruments
3821	yes	1.9	Laboratory Apparatus and Furniture
3823	yes	1.2	Industrial Instruments for Measurement, Display, and Control of Process Variables; and Related Products
3825	yes	4.2	Instruments for Measuring and Testing of Electricity and Electrical Signals
3826	yes	1.6	Laboratory Analytical Instruments
3827	yes	10.0	Optical Instruments and Lenses
3829	yes	7.2	Measuring and Controlling Devices, NEC
3841	yes	1.3	Surgical and Medical Instruments and Apparatus
3842	yes	2.1	Orthopedic, Prosthetic, and Surgical Appliances and Supplies
3843	yes	2.1	Dental Equipment and Supplies
3844	yes	1.3	X-Ray Apparatus and Tubes and Related Irradiation Apparatus
3999	yes	2.1	Manufacturing Industries, NEC

A.3.3 Additional results

Figure A.2: Spending response for durable and nondurable goods



Notes: The figure plots impulse response functions of military spending for the baseline sample estimated using specification (1.20). Shaded regions mark 80 percent confidence bands based on standard errors that are clustered at the industry level.

A.3.4 Robustness

Table A.7: Sectoral multipliers when controlling for Ramey's news variable

	Durable goods			Nondurable goods		
Years after shock	1	2	3	1	2	3
Gross output	0.26	0.23	0.73	2.83	2.49	2.83
Value added	0.44	0.38	0.71	0.97	0.83	1.13
Cost of materials	0.05	0.01	0.10	1.82	1.31	1.49
Energy expenditures	0.010	0.006	0.009	0.064	0.066	0.047
Employment (employees per year per $1m)	2.04	3.47	7.42	17.73	18.16	21.78

Notes: See notes of Table 1.4. The estimates are based on specification (1.20), augmented by Ramey's news variable interacted with industry indicators.

Table A.8: Sectoral multipliers when allowing for industry-specific effects of monetary policy

	Durable goods			Nondurable goods		
Years after shock	1	2	3	1	2	3
Gross output	0.34	0.21	0.45	2.58	2.67	3.24
Value added	0.40	0.35	0.55	0.97	0.94	1.29
Cost of materials	0.00	-0.10	-0.08	1.61	1.52	1.92
Energy expenditures	0.008	0.005	0.007	0.056	0.063	0.050
Employment (employees per year per $1m)	2.29	3.29	5.75	15.56	15.25	17.66

Notes: See notes of Table 1.4. The estimates are based on specification (1.20), augmented by interactions of the real interest rate with industry indicators.

Table A.9: Sectoral multipliers for a smoothing parameter of 400

	Durable goods			Nondurable goods		
Years after shock	1	2	3	1	2	3
Gross output	-0.05	-0.05	0.19	2.24	2.27	2.86
Value added	0.17	0.19	0.40	0.81	0.76	1.15
Cost of materials	-0.14	-0.20	-0.21	1.15	1.04	1.45
Energy expenditures	0.005	0.003	0.004	0.066	0.073	0.060
Employment (employees per year per $1m)	1.39	3.23	6.37	15.96	15.32	18.45

Notes: See notes of Table 1.4. The results in this table are obtained when the trends for all variables are extracted with a smoothing parameter of 400.

Table A.10: Sectoral multipliers for a smoothing parameter of 6000

	Durable goods			Nondurable goods		
Years after shock	1	2	3	1	2	3
Gross output	0.52	0.50	0.78	2.82	2.94	3.38
Value added	0.52	0.51	0.73	1.07	1.07	1.35
Cost of materials	0.11	0.08	0.13	1.88	1.86	2.18
Energy expenditures	0.010	0.007	0.009	0.050	0.051	0.038
Employment (employees per year per $1m)	2.25	3.34	5.65	15.19	14.77	16.40

Notes: See notes of Table 1.4. The results in this table are obtained when the trends for all variables are extracted with a smoothing parameter of 6000.

Table A.11: Sectoral multipliers estimated from 3-digit SIC industries

	Durable goods		
Years after shock	1	2	3
Gross output	-0.24	-0.23	0.02
Value added	-0.12	-0.06	0.01
Cost of materials	0.17	0.17	0.31
Energy expenditures	0.002	0.002	0.003
Employment (employees per year per $1m)	-2.67	-0.95	0.20

Notes: See notes of Table 1.4.

A.3.5 Impulse response functions of prices

I estimate the specification

$$\begin{aligned}\frac{P_{i,t+h}-P^T_{i,t-1}}{P^T_{i,t-1}} &= \alpha_h \frac{G_{i,t}-G^T_{i,t-1}}{VA^T_{i,t-1}} + \sum_{k=1}^{2}\beta_h^k \frac{P_{i,t-k}-P^T_{i,t-k-1}}{P^T_{i,t-k-1}} \\ &\quad + \sum_{k=1}^{2}\gamma_h^k \frac{G_{i,t-k}-G^T_{i,t-k-1}}{VA^T_{i,t-k-1}} + \delta_{i,h} + \zeta_{t,h} + \varepsilon_{i,t+h}, \qquad \text{(A.19)}\end{aligned}$$

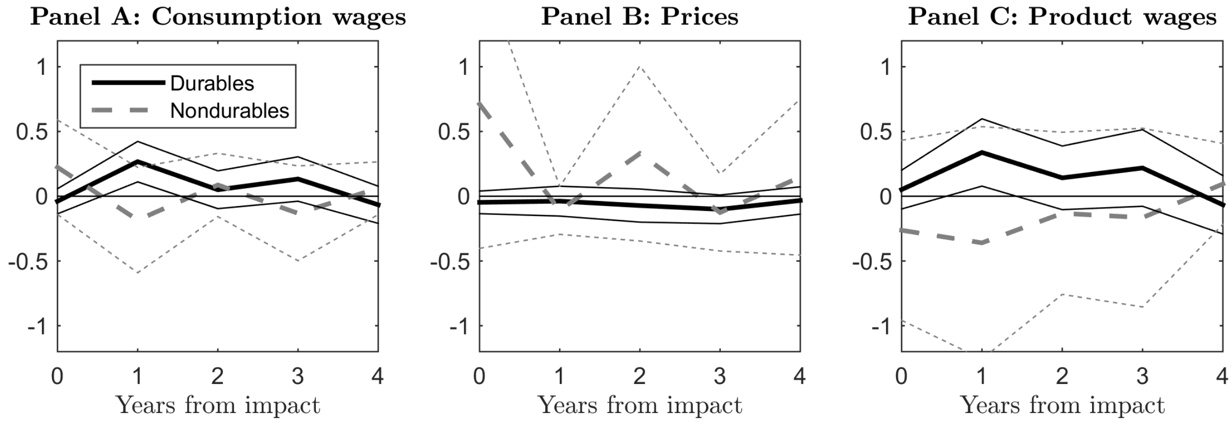
where $P_{i,t}$ denotes the industry's consumption wage, price, or product wage. The coefficients $\{\alpha_h\}_{h=0}^4$ are now akin to semi-elasticities and interpreted as follows. When the difference between government spending and its trend, normalized by the sector's value added, rises by one, the government engages, on average, in further spending in subsequent periods as shown in Panel A of Figure 1.3. This spending path is associated with a price response, expressed as a percentage deviation from trend, given by $\{\alpha_h\}_{h=0}^4$. As before, the spending variables on the right-hand side of equation (A.19) is instrumented with (1.21) and its two lags. The first stage Angrist-Pischke F-statistics are essentially identical to those reported in Table 1.3.

Figure A.3 displays the results. All prices, consumption wages (Panel A), product prices (Panel B), and product wages (Panel C) respond little in response to a fiscal shock. These effects are at odds with standard theory but not uncommon in the fiscal policy literature (see e.g. Nakamura and Steinsson, 2014).

There are a number of candidate explanations for the weak prices responses. First, due to their high price sensitivity, fluctuations of durable goods prices should generally be small and hard to detect.[3] Second, there may be composition effects. The military likely purchases a different basket of goods from a particular sector than the private sector. If the basket purchased by the military has a lower price than that purchased by the private sector, then

[3]See, for example, *House and Shapiro* (2008).

Figure A.3: Impulse response functions of prices



Notes: The figure plots impulse response functions estimated from specification (A.19) for the baseline sample as described in the text. The shock is a unit increase of government spending above trend, normalized by the sector's value added. Shaded regions mark 80 percent confidence bands based on standard errors that are clustered at the industry level.

no price response may be registred. Another possible explanation is that the government purchases goods in bulk and receives greater discounts than the private sector. Given that many nondurable goods sectors in the sample produce food, this explanation seems plausible. Yet, the weak responses of prices after fiscal shocks remain puzzling.

A.3.6 State dependence

Table A.12: First stages of the state-dependent specification

First stage dep. Variable		Horizon h	0	1	2	3	4
	AP F-statistic		132.8	116.1	125.3	115.8	38.2
$F_{i,t-1}\frac{G_{i,t}-G^T_{i,t-1}}{VA^T_{i,t-1}}$	p-value		0.00	0.00	0.00	0.00	0.00
	R-squared		0.44	0.45	0.49	0.50	0.51
	AP F-statistic		93.0	143.7	257.62	237.6	168.7
$(1-F_{i,t-1})\frac{G_{i,t}-G^T_{i,t-1}}{VA^T_{i,t-1}}$	p-value		0.00	0.00	0.00	0.00	0.00
	R-squared		0.43	0.54	0.57	0.58	0.59
	AP F-statistic		95.9	92.2	80.1	47.1	18.2
$F_{i,t-1}\frac{G_{i,t-1}-G^T_{i,t-2}}{VA^T_{i,t-2}}$	p-value		0.00	0.00	0.00	0.00	0.00
	R-squared		0.46	0.49	0.51	0.51	0.53
	AP F-statistic		205.3	244.4	204.6	161.4	96.5
$(1-F_{i,t-1})\frac{G_{i,t-1}-G^T_{i,t-2}}{VA^T_{i,t-2}}$	p-value		0.00	0.00	0.00	0.00	0.00
	R-squared		0.53	0.58	0.58	0.59	0.60
	AP F-statistic		171.4	186.9	198.1	103.6	59.5
$F_{i,t-1}\frac{G_{i,t-2}-G^T_{i,t-3}}{VA^T_{i,t-3}}$	p-value		0.00	0.00	0.00	0.00	0.00
	R-squared		0.45	0.47	0.49	0.50	0.51
	AP F-statistic		166.3	196.5	306.6	259.4	189.7
$(1-F_{i,t-1})\frac{G_{i,t-2}-G^T_{i,t-3}}{VA^T_{i,t-3}}$	p-value		0.00	0.00	0.00	0.00	0.00
	R-squared		0.60	0.60	0.60	0.61	0.61
Observations			2115	2039	1963	1887	1811

Notes: The table shows the first stages of the 2SLS estimator of specification (1.22) when the dependent variable is value added. For all other dependent variables, the F-statistics are virtually identical. The instruments are (1.21) and its two lags, interacted with $F_{i,t-1}$ and $1-F_{i,t-1}$. AP F-statistic stands for Angrist-Pischke F-statistic and the subsequent lines show the associated p-values.

APPENDIX B

Chapter 2 Appendices

B.1 Appendix: Proofs of the Propositions

Lemma 1 *Under the assumptions in Section 2.2, the unique competitive equilibrium of the model is characterized by equations (2.9) and (2.10), where* $\Phi_j = \sigma(1-\varrho_j)(1-\beta\varrho_j) - \kappa\varrho_j$ *for* $j = \{r, u, my, m\pi\}$ *are constants that are independent of monetary policy.*

Proof: Consider the system (2.1), (2.2) and (TR1) with the exogenous processes (2.3) to (2.6). Using (TR1) to eliminate i_t in (2.2) gives the system

$$\pi_t = \kappa y_t + \beta E_t[\pi_{t+1}] + u_t, \tag{B1}$$

$$y_t = E_t[y_{t+1}] - \frac{1}{\sigma}\left(\phi_\pi(\pi_t + m_t^\pi) + \phi_y(y_t + m_t^y) - r_t^e - E_t[\pi_{t+1}]\right). \tag{B2}$$

We conjecture an equilibrium solution

$$\pi_t = s_{\pi r} r_t^e + s_{\pi m^\pi} m_t^\pi + s_{\pi m^y} m_t^y + s_{\pi u} u_t$$

$$y_t = s_{yr} r_t^e + s_{ym^\pi} m_t^\pi + s_{ym^y} m_t^y + s_{yu} u_t$$

and solve for the eight unknown coefficients $s_{h,j}$ for $h = \pi, y$ and $j = r, u, my, m\pi$. Substituting the conjectured relationships into (B1) and (B2) and using (2.3) to (2.6) to evaluate the expectations gives

$$\begin{aligned} 0 &= (s_{\pi r} - \kappa s_{yr} - \beta s_{\pi r} \varrho_r) r_t^e + (s_{\pi m^\pi} - \kappa s_{ym^\pi} - \beta s_{\pi m^\pi} \varrho_{m^\pi}) m_t^\pi \\ &\quad + (s_{\pi m^y} - \kappa s_{ym^y} - \beta s_{\pi m^y} \varrho_{m^y}) m_t^y + (s_{\pi u} - \kappa s_{yu} - 1 - \beta s_{\pi u} \varrho_u) u_t \end{aligned}$$

and

$$\begin{aligned} 0 &= [(\varrho_r - 1) \sigma s_{yr} - \phi_\pi s_{\pi r} + 1 - \phi_y s_{yr} + s_{\pi r} \varrho_r] r_t^e \\ &\quad + [(\varrho_{m^\pi} - 1) \sigma s_{ym^\pi} - \phi_\pi s_{\pi m^\pi} - \phi_\pi - \phi_y s_{ym^\pi} + s_{\pi m^\pi} \varrho_{m^\pi}] m_t^\pi \\ &\quad + [(\varrho_{m^y} - 1) \sigma s_{ym^y} - \phi_\pi s_{\pi m^y} - \phi_y - \phi_y s_{ym^y} + s_{\pi m^y} \varrho_{m^y}] m_t^y \\ &\quad + [(\varrho_u - 1) \sigma s_{yu} - \phi_\pi s_{\pi u} - \phi_y s_{yu} + s_{\pi u} \varrho_u] u_t. \end{aligned}$$

Each coefficient in these expressions must be zero. This gives a system of eight equations in eight unknowns. Solving for these unknowns gives the coefficients in (2.9) and (2.10). (By collecting terms for each of the shocks r_t^e, m_t^π, m_t^y, and u_t one can split the system into four sub-systems, each with two equations and two unknowns. The four subsystems can then be solved separately.)

Uniqueness follows from the determinacy condition (2.7). ■

Proposition 1

(i) If the only shocks to the model are cost-push shocks, then the optimal policy requires that the Taylor rule coefficients lie on the affine manifold (2.11).

(ii) For any ϕ_π^* and ϕ_y^* satisfying (2.11), the equilibrium is

$$\pi_t = \frac{1-\beta\varrho_u}{\kappa^2+\alpha\left(1-\beta\varrho_u\right)^2}u_t, \qquad y_t = -\frac{\alpha\kappa}{\alpha\kappa^2+\left(1-\beta\varrho_u\right)^2}u_t. \tag{B3}$$

Proof: Substituting (2.9) and (2.10) (and using $r_t^e = m_t^y = m_t^\pi = 0$) into (2.8) gives

$$V\left[u_t\right]\left(\alpha\left(\frac{\phi_y+\left(1-\varrho_u\right)\sigma}{\Phi_u+\phi_y\left(1-\beta\varrho_u\right)+\kappa\phi_\pi}\right)^2+\left(\frac{\varrho_u-\phi_\pi}{\Phi_u+\phi_y\left(1-\beta\varrho_u\right)+\kappa\phi_\pi}\right)^2\right)$$

The first order condition for ϕ_π requires

$$0=\alpha\kappa\left(\phi_y+\left(1-\varrho_u\right)\sigma\right)+\left(\varrho_u-\phi_\pi\right)\left(1-\beta\varrho_u\right)$$

which implies (**??**). The first order condition for ϕ_y requires

$$0=\alpha\kappa\phi_y\left(\phi_\pi-\varrho_u\right)+\alpha\kappa\sigma\left(1-\varrho_u\right)\left(\phi_\pi-\varrho_u\right)-\left(\varrho_u-\phi_\pi\right)^2\left(1-\beta\varrho_u\right).$$

This condition can be satisfied either by setting $\phi_\pi=\varrho_u$ or by (2.11). This establishes (i).

To establish (ii), use $\Phi_u=\left(1-\varrho_u\right)\left(1-\beta\varrho_u\right)\sigma-\kappa\varrho_u$ to write the equilibrium conditions as

$$\pi_t = \frac{\phi_y+\left(1-\varrho_u\right)\sigma}{\left(1-\varrho_u\right)\left(1-\beta\varrho_u\right)\sigma+\kappa\left(\phi_\pi-\varrho_u\right)+\phi_y\left(1-\beta\varrho_u\right)}u_t,$$

$$y_t = \frac{\varrho_u-\phi_\pi}{\left(1-\varrho_u\right)\left(1-\beta\varrho_u\right)\sigma+\kappa\left(\phi_\pi-\varrho_u\right)+\phi_y\left(1-\beta\varrho_u\right)}u_t.$$

Substituting (2.11) gives (B3). This establishes (ii). ■

Proposition 2 *Suppose r_t^e and u_t are i.i.d. over time and have covariance $Cov\left[r_t^e,u_t\right]$. The Taylor rule coefficients satisfying (2.11') and $\phi_y\to\infty$ are optimal.*

Proof: When r_t^e and u_t are i.i.d. equations (2.9) and (2.10) simplify to

$$\pi_t = \frac{\kappa}{\sigma + \phi_y + \kappa\phi_\pi} r_t^e + \frac{\phi_y + \sigma}{\sigma + \phi_y + \kappa\phi_\pi} u_t$$

$$y_t = \frac{1}{\sigma + \phi_y + \kappa\phi_\pi} r_t^e - \frac{\phi_\pi}{\sigma + \phi_y + \kappa\phi_\pi} u_t$$

Plugging them into objective (2.8), and simplifying, shows that the optimal ϕ_y and ϕ_π minimize

$$(\sigma + \phi_y + \kappa\phi_\pi)^{-2} \left[\left(1 + \alpha\kappa^2\right) V\left[r^e\right] + \left(\alpha\left(\phi_y + \sigma\right)^2 + \left(\phi_\pi\right)^2\right) V\left[u\right] + 2\left(\alpha\kappa\left(\phi_y + \sigma\right) - \phi_\pi\right) Cov\left[r_t^e, u_t\right] \right]$$

First we fix the overall strength of the policy response by setting $\sigma + \phi_y + \kappa\phi_\pi = A$ and minimizing

$$A^{-2} \left\{ \left(1 + \alpha\kappa^2\right) V\left[r^e\right] + \left(\alpha\left(\phi_y + \sigma\right)^2 + \left(\phi_\pi\right)^2\right) V\left[u\right] + 2\left(\alpha\kappa\left(\phi_y + \sigma\right) - \phi_\pi\right) Cov\left[r_t^e, u_t\right] \right\}$$

subject to $\sigma + \phi_y + \kappa\phi_\pi = A$.

The Lagrangian is

$$\begin{aligned} \mathcal{L} \ = \ & A^{-2} \left\{ \left(1 + \alpha\kappa^2\right) V\left[r^e\right] + \left(\alpha\left(\phi_y + \sigma\right)^2 + \left(\phi_\pi\right)^2\right) V\left[u\right] + 2\left(\alpha\kappa\left(\phi_y + \sigma\right) - \phi_\pi\right) Cov\left[r_t^e, u_t\right] \right\} + \\ & \lambda\left[A - \sigma - \phi_y - \kappa\phi_\pi\right] \end{aligned}$$

and the first order conditions w.r.t. ϕ_y and ϕ_π, respectively, are

$$A^{-2}\left\{2\alpha\left(\phi_y + \sigma\right) V\left[u\right] + 2\alpha\kappa Cov\left[r_t^e, u_t\right]\right\} = \lambda$$

$$A^{-2}\left\{2\left(\phi_\pi\right) V\left[u\right] - 2Cov\left[r_t^e, u_t\right]\right\} = \kappa\lambda$$

Combining them yields (2.11').

Since the objective is decreasing in A, it is optimal to let ϕ_y approach infinity. ■

Proposition 3 *Suppose all shocks are white noise, that is,* $\varrho_r = \varrho_u = \varrho_{my} = \varrho_{m\pi} = 0$. *Then the minimization of (2.8) subject to (2.9) and (2.10) yields the optimal Taylor rule coefficients given by (2.12) and (2.13).*

Proof: Setting $\varrho_r = \varrho_u = \varrho_{my} = \varrho_{m\pi} = 0$ in (2.9) and (2.10) and substituting into the objective (2.8) and using the fact that the shocks are (by assumption) uncorrelated implies that the central bank wants to choose parameters ϕ_y and ϕ_π to minimize

$$
\begin{aligned}
&(\sigma + \phi_y + \kappa\phi_\pi)^{-2} \\
&\left(\left[\alpha\kappa^2 + 1\right] V\left[r_t^n\right] + \left[\alpha\kappa^2 + 1\right](\phi_\pi)^2 V\left[m_t^\pi\right] + \left[\alpha\kappa^2 + 1\right](\phi_y)^2 V\left[m_t^y\right] + \left[\alpha(\phi_y + \sigma)^2 + (\phi_\pi)^2\right] V\left[u_t\right]\right)
\end{aligned}
$$

The first order condition for ϕ_y is

$$
\begin{aligned}
&\left\{\left[\alpha\kappa^2 + 1\right] V\left[m_t^y\right]\phi_y + \alpha(\phi_y + \sigma) V\left[u_t\right]\right\}(\sigma + \phi_y + \kappa\phi_\pi) \qquad \text{(B.1)} \\
= \; &\left[\alpha\kappa^2 + 1\right] V\left[r_t^n\right] + \left[\alpha\kappa^2 + 1\right] V\left[m_t^\pi\right](\phi_\pi)^2 + \left[\alpha\kappa^2 + 1\right] V\left[m_t^y\right](\phi_y)^2 + \left[\alpha(\phi_y + \sigma)^2 + (\phi_\pi)^2\right] V\left[u_t\right]
\end{aligned}
$$

The first order condition w.r.t. ϕ_π is

$$
\begin{aligned}
&\frac{\phi_\pi}{\kappa}\left\{\left[\alpha\kappa^2 + 1\right] V\left[m_t^\pi\right] + V\left[u_t\right]\right\}(\sigma + \phi_y + \kappa\phi_\pi) \qquad \text{(B.2)} \\
= \; &\left[\alpha\kappa^2 + 1\right] V\left[r_t^n\right] + \left[\alpha\kappa^2 + 1\right] V\left[m_t^\pi\right](\phi_\pi)^2 + \left[\alpha\kappa^2 + 1\right] V\left[m_t^y\right](\phi_y)^2 + \left[\alpha(\phi_y + \sigma)^2 + (\phi_\pi)^2\right] V\left[u_t\right]
\end{aligned}
$$

It is immediate to see that this implies

$$
\phi_\pi = \kappa\frac{\left[\alpha\kappa^2 + 1\right] V\left[m_t^y\right] + \alpha V\left[u_t\right]}{\left[\alpha\kappa^2 + 1\right] V\left[m_t^\pi\right] + V\left[u_t\right]}\phi_y + \frac{\alpha\kappa\sigma V\left[u_t\right]}{\left[\alpha\kappa^2 + 1\right] V\left[m_t^\pi\right] + V\left[u_t\right]} \qquad \text{(B6)}
$$

We can rewrite equation (B.2) as

$$
\begin{aligned}
& \kappa\left[\alpha\kappa^2+1\right]V\left[r_t^n\right]+\kappa\left[\alpha\kappa^2+1\right]V\left[m_t^y\right](\phi_y)^2 \\
= \; & \phi_\pi\left[\alpha\kappa^2+1\right](\sigma+\phi_y+\kappa\phi_\pi)V\left[m_t^\pi\right]-\kappa\left[\alpha\kappa^2+1\right]V\left[m_t^\pi\right](\phi_\pi)^2+ \\
& \phi_\pi(\sigma+\phi_y+\kappa\phi_\pi)V\left[u_t\right]-\kappa\left[\alpha(\phi_y+\sigma)^2+(\phi_\pi)^2\right]V\left[u_t\right]
\end{aligned}
$$

Cancelling like terms and simplifying gives

$$
\begin{aligned}
& \frac{\kappa}{\sigma+\phi_y}\left[\alpha\kappa^2+1\right]V\left[r_t^n\right]+\frac{\kappa}{\sigma+\phi_y}\left[\alpha\kappa^2+1\right]V\left[m_t^y\right](\phi_y)^2 \\
= \; & \phi_\pi\left\{\left[\alpha\kappa^2+1\right]V\left[m_t^\pi\right]+V\left[u_t\right]\right\}-\kappa\alpha\phi_y V\left[u_t\right]-\kappa\alpha\sigma V\left[u_t\right]
\end{aligned}
$$

Using condition (B6) we have

$$
\phi_\pi\left(\left[\alpha\kappa^2+1\right]V\left[m_t^\pi\right]+V\left[u_t\right]\right)-\alpha\kappa\sigma V\left[u_t\right]=\kappa\left(\left[\alpha\kappa^2+1\right]V\left[m_t^y\right]+\alpha V\left[u_t\right]\right)\phi_y
$$

Eliminating this term gives

$$
\begin{aligned}
& \frac{\kappa}{\sigma+\phi_y}\left[\alpha\kappa^2+1\right]V\left[r_t^n\right]+\frac{\kappa}{\sigma+\phi_y}\left[\alpha\kappa^2+1\right]V\left[m_t^y\right](\phi_y)^2 \\
= \; & \kappa\left(\left[\alpha\kappa^2+1\right]V\left[m_t^y\right]+\alpha V\left[u_t\right]\right)\phi_y-\kappa\alpha\phi_y V\left[u_t\right].
\end{aligned}
$$

Finally, we cancel terms to get (2.12). To find (2.13) use condition (B6) and rearrange terms.

■

Model with signal extraction To solve the model with signal extraction, we closely follow the setup in Svensson and Woodford (2003, 2004). The proofs of the results use the following notation and calculations.

The model can be written as

$$\begin{pmatrix} X_{t+1} \\ \tilde{E}\mathbb{E}_t\left[x_{t+1}\right] \end{pmatrix} = A \begin{pmatrix} X_t \\ x_t \end{pmatrix} + B\left(i_t - \rho\right) + \begin{pmatrix} s_{t+1} \\ 0 \end{pmatrix} \tag{B7}$$

where $X_t = (r_t^e, u_t, m_t^y, m_t^\pi)'$; $s_{t+1} = \left(\varepsilon_{t+1}^r, \varepsilon_{t+1}^u, \varepsilon_{t+1}^y, \varepsilon_{t+1}^\pi\right)'$; $x_t = (y_t, \pi_t)'$; and $\tilde{E} = \begin{pmatrix} 0 & 1 \\ \sigma & 1 \end{pmatrix}$. Partition the matrices A and B as follows

$$A = \begin{pmatrix} A_1 \\ A_2 \end{pmatrix} = \begin{pmatrix} A_{11} & A_{12} \\ A_{21} & A_{22} \end{pmatrix}, B = \begin{pmatrix} B_1 \\ B_2 \end{pmatrix}$$

and let

$$A_{11} = \begin{pmatrix} \varrho_r & 0 & 0 & 0 \\ 0 & \varrho_u & 0 & 0 \\ 0 & 0 & \varrho_{m^y} & 0 \\ 0 & 0 & 0 & \varrho_{m^\pi} \end{pmatrix}, A_{12} = \begin{pmatrix} 0 & 0 \\ 0 & 0 \\ 0 & 0 \\ 0 & 0 \end{pmatrix}, B_1 = \begin{pmatrix} 0 \\ 0 \\ 0 \\ 0 \end{pmatrix}$$

$$A_{21} = \begin{pmatrix} 0 & -\frac{1}{\beta} & 0 & 0 \\ -1 & 0 & 0 & 0 \end{pmatrix}, A_{22} = \begin{pmatrix} -\frac{\kappa}{\beta} & \frac{1}{\beta} \\ \sigma & 0 \end{pmatrix} \; B_2 = \begin{pmatrix} 0 \\ 1 \end{pmatrix}$$

The flow objective is

$$y_t^2 + \alpha\pi_t^2 = x_t' W x_t, \; W = \begin{pmatrix} 1 & 0 \\ 0 & \alpha \end{pmatrix} \tag{B8}$$

and the observation equations are

$$Z_t = \begin{pmatrix} y_t^m \\ \\ \pi_t^m \end{pmatrix} = D \begin{pmatrix} X_t \\ \\ x_t \end{pmatrix}, \text{ with } D = (D_1, D_2)\,,\; D_1 = \begin{pmatrix} 0 & 0 & 1 & 0 \\ \\ 0 & 0 & 0 & 1 \end{pmatrix} \text{ and } D_2 = \begin{pmatrix} 1 & 0 \\ \\ 0 & 1 \end{pmatrix} \tag{B9}$$

Notice that the measurement errors are included in X_t. This allows for arbitrary persistence of both types of measurement error.

The central bank's information set is $I_t^{CB} = \left\{\Theta, y_{t-j}^m, \pi_{t-j}^m : j \geq 0\right\}$ where Θ is a vector of all model parameters.

We write the Taylor rule as

$$i_t - \rho = F x_{t|t} \tag{B10}$$

where $F = (\psi_y, \psi_\pi)$. Notice the difference to Svensson and Woodford who write the policy as a linear function of the estimates of the state variables X_t.

Next, we conjecture that

$$x_{t|t} = G X_{t|t} \tag{B11}$$

Then the upper block of (B7) gives

$$X_{t+1} = A_{11} X_t + s_{t+1}$$

and taking expectations (based on I_t^{CB}) yields

$$X_{t+1|t} = A_{11} X_{t|t}.$$

Taking expectations of the lower block of (B7) gives

$$\tilde{E} x_{t+1|t} = A_{21} X_{t|t} + A_{22} x_{t|t} + B_2 \left(i_t - \rho\right).$$

Combining these equations with (B10) and (B11), we arrive at

$$x_{t|t} = [A_{22}]^{-1} \left(-A_{21} + \tilde{E}GA_{11} - B_2FG \right) X_{t|t}.$$

Hence, G must satisfy

$$G = [A_{22}]^{-1} \left(-A_{21} + \tilde{E}GA_{11} - B_2FG \right). \tag{B12}$$

Next we conjecture that

$$x_t = G^1 X_t + \left(G - G^1 \right) X_{t|t} \tag{B13}$$

and rewrite the observation equation (B9) as

$$\begin{aligned} Z_t &= D_1 X_t + D_2 x_t \\ &= \left(D_1 + D_2 G^1 \right) X_t + D_2 \left(G - G^1 \right) X_{t|t} \end{aligned}$$

Following Svensson and Woodford,

$$Z_t = LX_t + MX_{t|t} \tag{B14}$$

where

$$L = \left(D_1 + D_2 G^1 \right) \tag{B15}$$

and

$$M = D_2 \left(G - G^1 \right).$$

The state equation of the Kalman filter is

$$X_{t+1} = A_{11} X_t + s_{t+1}$$

and the observation equation is (B14). Svensson and Woodford (2004) show that the Kalman filter updating equation takes the form

$$X_{t|t} = X_{t|t-1} + KL\left(X_t - X_{t|t-1}\right) \tag{B16}$$

(their equation 26 with $v_t = 0$) and that it is possible to write

$$X_{t+1|t+1} = (I + KM)^{-1}\left[(I - KL)\,A_{11}X_{t|t} + KZ_{t+1}\right] \tag{B17}$$

(their equation 30) where

$$K = PL'\left(LPL'\right)^{-1}. \tag{B18}$$

Furthermore,

$$P = \mathbb{E}\left[\left(X_t - X_{t|t-1}\right)\left(X_t - X_{t|t-1}\right)'\right] = A_{11}\left[P - PL'\left(LPL'\right)^{-1}LP\right]A_{11}' + \Sigma_s \tag{B19}$$

where Σ_s is the covariance matrix of the errors s_{t+1}.

Finally, one needs to find G^1. Again, following Svensson and Woodford (2003, 2004) we obtain

$$G^1 = [A_{22}]^{-1}\left(-A_{21} + \tilde{E}GKLA_{11} + \tilde{E}G^1(I - KL)\,A_{11}\right). \tag{B20}$$

Lemma 2 *The central bank's estimates of inflation and the output gap satisfy (2.14) and (2.15) where Φ_r and Φ_u are defined as in Lemma 1.*

Proof: Take conditional expectations based on the central bank's information set I_t^{CB} of equations (2.1) to (2.4) to obtain

$$\pi_t = \kappa y_t + \beta\pi_{t+1|t} + u_{t|t}$$

$$y_t = y_{t+1|t} - \frac{1}{\sigma}\left(i_t - \rho - r^e_{t|t} - \pi_{t+1|t}\right)$$

$$r^e_{t+1|t} = \varrho_r r^e_{t|t} \text{ , and } u_{t+1|t} = \varrho_u u_{t|t}$$

Imposing the Taylor rule (TR2) and repeating steps in the proof of Lemma 1 yields the desired result. ■

Lemma A1 *The welfare objective (2.8) can be decomposed as follows*

$$(1-\beta)\,\mathbb{E}\left[\sum_{t=0}^{\infty}\beta^t\left(y_t^2+\alpha\pi_t^2\right)\right] = \left(\alpha\mathbb{E}\left[\pi^2_{t|t}\right]+\mathbb{E}\left[y^2_{t|t}\right]\right)+\left(\alpha\mathbb{E}\left[\left(\pi_t-\pi_{t|t}\right)^2\right]+\mathbb{E}\left[\left(y_t-y_{t|t}\right)^2\right]\right)$$

Proof: See Svensson and Woodford (2003, 2004). ■

Lemma A2 $\alpha\mathbb{E}\left[\left(\pi_t-\pi_{t|t}\right)^2\right]+\mathbb{E}\left[\left(y_t-y_{t|t}\right)^2\right] = tr\left[P\left(G^1\left(I-KL\right)\right)'WG^1\left(I-KL\right)\right]$ *where* I *is the identiy matrix.*

Proof: Define $T = \alpha\mathbb{E}\left[\left(\pi_t-\pi_{t|t}\right)^2\right]+\mathbb{E}\left[\left(y_t-y_{t|t}\right)^2\right]$ and notice that $T = \mathbb{E}\left[\left(x_t-x_{t|t}\right)'W\left(x_t-x_{t|t}\right)\right]$ where W is defined in (B8) above. Equation (B13) implies that

$$x_t - x_{t|t} = G^1X_t + \left(G-G^1\right)X_{t|t} - GX_{t|t} = G^1\left(X_t - X_{t|t}\right)$$

so we obtain

$$T = \mathbb{E}\left[\left(X_t-X_{t|t}\right)'\left(G^1\right)'WG^1\left(X_t-X_{t|t}\right)\right].$$

Next subtract $X_{t|t}$ from X_t and use (B16) to get

$$\begin{aligned} X_t - X_{t|t} &= X_t - \left(X_{t|t-1} + KL\left(X_t - X_{t|t-1}\right)\right) \\ &= \left(I-KL\right)\left(X_t - X_{t|t-1}\right) \end{aligned}$$

Then

$$
\begin{aligned}
T &= \mathbb{E}\left[\left(X_t - X_{t|t-1}\right)'(I-KL)'\left(G^1\right)' W G^1 (I-KL)\left(X_t - X_{t|t-1}\right)\right] \\
&= \mathbb{E}\left[tr\left[\left(X_t - X_{t|t-1}\right)\left(X_t - X_{t|t-1}\right)'(I-KL)'\left(G^1\right)' W G^1 (I-KL)\right]\right] \\
&= tr\left[\mathbb{E}\left[\left(X_t - X_{t|t-1}\right)\left(X_t - X_{t|t-1}\right)'\right]\left(G^1 (I-KL)\right)' W G^1 (I-KL)\right]
\end{aligned}
$$

where $tr(\cdot)$ is the trace operator. Using (B19) yields the desired result. ■

Lemma A3 *Suppose all shocks are contemporaneously uncorrelated with each other and i.i.d over time. Then the term* $\alpha\mathbb{E}\left[\left(\pi_t - \pi_{t|t}\right)^2\right] + \mathbb{E}\left[\left(y_t - y_{t|t}\right)^2\right]$ *is independent of the Taylor rule coefficients.*

Proof: Given Lemma A2, it is sufficient to show that T is independent of the Taylor rule coefficients. If all shocks are i.i.d. over time then $A_{11} = 0_{4\times 4}$ and G^1 as defined in (B20) reduces to

$$G^1 = -\left[A_{22}\right]^{-1} A_{21}$$

which is independent of policy. Furthermore, $P = \Sigma_s$ (see B19) and because G^1 is independent of policy, so is L (defined in B15). Then $K = PL'\left(LPL'\right)^{-1}$is independent of policy (equation B18). As a result, T is independent of policy. ■

Lemma A4 *Suppose all shocks are contemporaneously uncorrelated with each other and i.i.d over time. Then* $r^e_{t|t}$ *and* $u_{t|t}$ *are independent of the Taylor rule coefficients.*

Proof: In the iid case with $A_{11} = 0_{4\times 4}$, equation (B17) reduces to

$$X_{t+1|t+1} = (I + KM)^{-1} KZ_{t+1}$$

Combining this with (B14) gives

$$X_{t|t} = (I + KM)^{-1} KZ_t = (I + KM)^{-1} K \left(LX_t + MX_{t|t}\right).$$

Rearranging yields

$$X_{t|t} = KLX_t.$$

In the proof of Lemma A3 we showed that K and L are independent of policy when shocks are i.i.d. Hence $r^e_{t|t}$ and $u_{t|t}$ are independent of the Taylor rule coefficients. ■

Proposition 4 *Suppose all shocks are contemporaneously uncorrelated with each other and i.i.d over time. Then the coefficients* $\left\{\psi^*_y, \psi^*_\pi\right\}$ *satisfying (11") and* $\psi_y \to \infty$ *are optimal.*

Proof: Lemmas A1 and A3 imply that minimizing objective (2.8) is equivalent to minimizing $\alpha\mathbb{E}\left[\pi^2_{t|t}\right] + \mathbb{E}\left[y^2_{t|t}\right]$. Furthermore, equations (2.14) and (2.15) simplify to

$$\pi_{t|t} = \frac{\kappa}{\sigma + \psi_y + \kappa\psi_\pi} r^e_{t|t} + \frac{\psi_y + \sigma}{\sigma + \psi_y + \kappa\psi_\pi} u_{t|t}$$

$$y_{t|t} = \frac{1}{\sigma + \psi_y + \kappa\psi_\pi} r^e_{t|t} - \frac{\psi_\pi}{\sigma + \psi_y + \kappa\psi_\pi} u_{t|t}$$

in the i.i.d. case. Lemma A4 shows that $r^e_{t|t}$ and $u_{t|t}$ are independent of the Taylor rule coefficients, but importantly, their covariance need not to be zero. The remainder of the proof is identical to the proof of Proposition 2. ■

Proposition 5 *Suppose the model is given by equations (2.1) to (2.6), (TR2), and the observation equations* $\pi^m_t = \pi_t + m^\pi_t$ *and* $y^m_t = y_t + m^y_t$*. The central bank uses the Kalman filter with information set* I^{CB}_t *to estimate the true state of the economy. Suppose further that* $V[u_t] = 0$ *and that* $\varrho_{my} = \varrho_{m\pi} = 0$*. Then, for any Taylor rule (TR2), with coefficients* ψ_y *and* ψ_π*, there exist coefficients* $\left\{\tilde{\phi}_y, \tilde{\phi}_\pi, \tilde{\nu}\right\}$ *such that the policy rule (TR3) generates the* *same equilibrium paths for all variables.*

Proof: Under the assumption that $V[u_t] = 0$ and $\varrho_{my} = \varrho_{m\pi} = 0$, it is possible to write the model as follows

$$X_t = r_t^e,\ s_{t+1} = \varepsilon_{t+1}^r,\ x_t = \begin{pmatrix} y_t \\ \pi_t \end{pmatrix},\ \tilde{E} = \begin{pmatrix} 0 & 1 \\ \sigma & 1 \end{pmatrix},\ A = \begin{pmatrix} A_{11} & A_{12} \\ A_{21} & A_{22} \end{pmatrix}, B = \begin{pmatrix} B_1 \\ B_2 \end{pmatrix}$$

$$A_{11} = \varrho_r,\ A_{12} = \begin{pmatrix} 0 & 0 \end{pmatrix}, B_1 = 0,\ A_{21} = \begin{pmatrix} 0 \\ -1 \end{pmatrix},\ A_{22} = \begin{pmatrix} -\frac{\kappa}{\beta} & \frac{1}{\beta} \\ \sigma & 0 \end{pmatrix}\ B_2 = \begin{pmatrix} 0 \\ 1 \end{pmatrix}$$

The observation equations are

$$Z_t = \begin{pmatrix} y_t^m \\ \pi_t^m \end{pmatrix} + v_t = D \begin{bmatrix} X_t \\ x_t \end{bmatrix} + v_t,\ D = \begin{pmatrix} 0 & 1 & 0 \\ 0 & 0 & 1 \end{pmatrix}$$

with $v_t = (m_t^y, m_t^\pi)'$. Furthermore, $D_1 = (0,0)'$ and $D_2 = I_2$ Notice that A_{11} is a scalar. Because v_t is nonzero, the equations (and matrices) associated with the Kalman filter change somewhat (see Svensson and Woodford 2004) though (B17) continues to hold. Since A_{11} is a scalar, KL and KM are also scalars.

Using (B10) and (B11), write the interest rate as

$$i_t - \rho = (\psi_y, \psi_\pi)\left(y_{t|t}, \pi_{t|t}\right)' = Fx_{t|t} = FGX_{t|t}.$$

Using (B17) and substituting backwards yields

$$i_t = \sum_{j=0}^{\infty} FG\left(I + KM\right)^{-1}\left[\left(I - KL\right)A_{11}\left(I + KM\right)^{-1}\right]^j KZ_{t-j}$$

Since KL, and KM are scalars, set $\tilde{\nu} = (I - KL)\, A_{11}\, (I + KM)^{-1}$. Then write

$$i_t = \sum_{j=0}^{\infty} FG\,(I + KM)^{-1}\,\tilde{\nu}^j K Z_{t-j} = FG\,(I + KM)^{-1}\,KZ_t + \tilde{\nu} i_{t-1} = \tilde{\phi}_y y_t^m + \tilde{\phi}_\pi \pi_t^m + \tilde{\nu} i_{t-1}$$

where $\left(\tilde{\phi}_y, \tilde{\phi}_\pi\right) = FG\,(I + KM)^{-1}\,K.$ $\blacksquare$

APPENDIX C

Chapter 3 Appendices

C.1 Chapter 2: Basic Theory Appendix

C.1.1 Proof of Result 1

Suppose that the firm solves

$$\max p_x x - p_D F_D - p_M IM$$

subject to

$$x = \left[(1-\mu)^{\frac{1}{\psi}} \left[F_D\right]^{\frac{\psi-1}{\psi}} + \mu^{\frac{1}{\psi}} \left[IM\right]^{\frac{\psi-1}{\psi}} \right]^{\frac{\psi}{\psi-1}}$$

and

$$p_x = \left(\frac{Y}{x}\right)^{\frac{1}{\varepsilon}}$$

The first order conditions are

$$\left(1-\frac{1}{\varepsilon}\right)(Y)^{\frac{1}{\varepsilon}}(x)^{\frac{1}{\psi}-\frac{1}{\varepsilon}}(1-\mu)^{\frac{1}{\psi}}\left[F_D\right]^{-\frac{1}{\psi}} = p_D$$

$$\left(1-\frac{1}{\varepsilon}\right)(Y)^{\frac{1}{\varepsilon}}(x)^{\frac{1}{\psi}-\frac{1}{\varepsilon}}\mu^{\frac{1}{\psi}}[IM]^{-\frac{1}{\psi}}=p_M$$

Dividing one by the other gives

$$\frac{F_D^*}{IM^*}=\frac{1-\mu}{\mu}\left(\frac{p_M}{p_D}\right)^{\psi}.$$

The same equation can be obtained under perfect competition.

Now take the production function and multiply it by p_x

$$p_x x=p_x\left[(1-\mu)^{\frac{1}{\psi}}[F_D]^{\frac{\psi-1}{\psi}}+(p_M)^{-\frac{\psi-1}{\psi}}\mu^{\frac{1}{\psi}}[p_M IM]^{\frac{\psi-1}{\psi}}\right]^{\frac{\psi}{\psi-1}}$$

Taking logs gives

$$\ln(p_x x)=\frac{\psi}{\psi-1}\ln\left(p_x\left[(1-\mu)^{\frac{1}{\psi}}[F_D]^{\frac{\psi-1}{\psi}}+(p_M)^{-\frac{\psi-1}{\psi}}\mu^{\frac{1}{\psi}}[p_M IM]^{\frac{\psi-1}{\psi}}\right]\right) \tag{C.1}$$

$$=\frac{\psi}{\psi-1}\ln\left(p_x\left[(1-\mu)^{\frac{1}{\psi}}\exp\left(\frac{\psi-1}{\psi}\ln[F_D]\right)+(p_M)^{-\frac{\psi-1}{\psi}}\mu^{\frac{1}{\psi}}\exp\left(\frac{\psi-1}{\psi}\ln[p_M IM]\right)\right]\right) \tag{C.2}$$

Before differentiating, recall the assumption that the firm takes prices p_M as given and that it cannot change p_x after learning about the shock. Then

$$\frac{\partial\ln p_x x}{\partial\ln p_M M}=\frac{\psi}{\psi-1}\frac{p_x(P_M)^{-\frac{\psi-1}{\psi}}\mu^{\frac{1}{\psi}}\exp\left(\frac{\psi-1}{\psi}\ln[p_M IM]\right)\frac{\psi-1}{\psi}}{p_x\left[(1-\mu)^{\frac{1}{\psi}}\exp\left(\frac{\psi-1}{\psi}\ln[F_D]\right)+(p_M)^{-\frac{\psi-1}{\psi}}\mu^{\frac{1}{\psi}}\exp\left(\frac{\psi-1}{\psi}\ln[p_M IM]\right)\right]} \tag{C.3}$$

$$=\frac{1}{1+\left(\frac{1-\mu}{\mu}\right)^{\frac{1}{\psi}}\left[\frac{F_D}{IM}\right]^{\frac{\psi-1}{\psi}}} \tag{C.4}$$

We evaluate this elasticity at

$$\frac{F_D^*}{IM} = \frac{IM^*}{IM}\frac{1-\mu}{\mu}\left(\frac{p_M}{p_D}\right)^{\psi}$$

so that

$$\frac{\partial \ln p_x x}{\partial \ln p_M IM} = \frac{1}{1+\left(\frac{IM^*}{IM}\right)^{\frac{\psi-1}{\psi}}\frac{1-\mu}{\mu}\left(\frac{p_M}{p_D}\right)^{\psi-1}}$$

C.1.2 On Flexibility in Domestic Inputs

Under the assumption of perfect competition, the first order conditions are:

$$x\left(1-\mu\right) = \left(p_D\right)^{\psi} F_D$$

$$x\mu = \left(p_M\right)^{\psi} IM$$

If the firm takes prices p_x, p_M, and p_D as given, the following elasticities are immediate:

$$\frac{\partial \ln\left(p_x x\right)}{\partial \ln\left(p_D F_D\right)} = \frac{\partial \ln\left(p_x x\right)}{\partial \ln\left(p_M M\right)} = \frac{\partial \ln\left(p_D F_D\right)}{\partial \ln\left(p_M M\right)} = 1.$$

The above equations demonstrate that a constant returns to scale production function combined with these assumptions on market structure imply that the output elasticity will equal one for all values of the elasticity of substitution. For this reason, we require some assumptions limiting the flexibility of domestic inputs following the import disruption.

Below we show an alternative way of understanding the interaction of competitive factor markets, changes in domestic inputs, and the mapping of the output elasticity into parameter values for the elasticity of substitution. Consider the total derivative of $\ln(x)$:

$$d\ln x = \frac{\partial \ln x}{\partial IM} d\ln IM + \frac{\partial \ln x}{\partial F} d\ln F \tag{C.5}$$

$$d\ln x = \frac{\mu^{\frac{1}{\psi}}(IM)^{\frac{\psi-1}{\psi}} d\ln IM}{(1-\mu)^{\frac{1}{\psi}}[F_D]^{\frac{\psi-1}{\psi}} + \mu^{\frac{1}{\psi}}[IM]^{\frac{\psi-1}{\psi}}} + \frac{(1-\mu)^{\frac{1}{\psi}}(F_D)^{\frac{\psi-1}{\psi}} d\ln F_D}{(1-\mu)^{\frac{1}{\psi}}[F_D]^{\frac{\psi-1}{\psi}} + \mu^{\frac{1}{\psi}}[IM]^{\frac{\psi-1}{\psi}}} \tag{C.6}$$

Dividing by $d\ln IM$ yields:

$$\frac{d\ln x}{d\ln IM} = \frac{\mu^{\frac{1}{\psi}}(IM)^{\frac{\psi-1}{\psi}}}{(1-\mu)^{\frac{1}{\psi}}[F_D]^{\frac{\psi-1}{\psi}} + \mu^{\frac{1}{\psi}}[IM]^{\frac{\psi-1}{\psi}}} + \frac{(1-\mu)^{\frac{1}{\psi}}(F_D)^{\frac{\psi-1}{\psi}}}{(1-\mu)^{\frac{1}{\psi}}[F_D]^{\frac{\psi-1}{\psi}} + \mu^{\frac{1}{\psi}}[IM]^{\frac{\psi-1}{\psi}}} \frac{d\ln F_D}{d\ln IM}$$

Now, as before, combining the first order conditions from the profit maximization problem, we have:

$$\frac{F_D(\cdot)}{IM} = \frac{1-\mu}{\mu}\left(\frac{p_D}{p_M}\right)^{-\psi} \tag{C.7}$$

Log-differentiating this expression:

$$\begin{aligned} d\ln\left(\frac{F_D}{IM}\right) &= -\psi d\ln\left(\frac{p_D}{p_M}\right) \\ d\ln F_D - d\ln IM &= -\psi d\ln\left(\frac{p_D}{p_M}\right) \\ \frac{d\ln F_D}{d\ln IM} &= 1 - \psi\frac{d\ln\left(\frac{p_D}{p_M}\right)}{d\ln IM} \end{aligned} \tag{C.8}$$

Finally, we have:

$$\frac{d\ln x}{d\ln IM} = \frac{\mu^{\frac{1}{\psi}}(IM)^{\frac{\psi-1}{\psi}}}{(1-\mu)^{\frac{1}{\psi}}[F_D]^{\frac{\psi-1}{\psi}} + \mu^{\frac{1}{\psi}}[IM]^{\frac{\psi-1}{\psi}}} + \frac{(1-\mu)^{\frac{1}{\psi}}(F_D)^{\frac{\psi-1}{\psi}}\left[1 - \psi\frac{d\ln\left(\frac{p_D}{p_M}\right)}{d\ln IM}\right]}{(1-\mu)^{\frac{1}{\psi}}[F_D]^{\frac{\psi-1}{\psi}} + \mu^{\frac{1}{\psi}}[IM]^{\frac{\psi-1}{\psi}}} \tag{C.9}$$

Thus, if there is no change in the relative input price following the disruption in IM of the firm: $\frac{d\ln\left(\frac{p_M}{p_D}\right)}{d\ln IM} = 0$, then the output elasticity will be equal to one regardless of the value of ψ. On the other hand, any assumptions that yield a non-zero change in the relative input prices will then yield the result that $\frac{d\ln x}{d\ln IM} = 1$ provided $\psi \to 0$.

C.2 Chapter 3: Data Appendix

C.2.1 Matching Corporate Directories to the Business Register

The discussion below is an abbreviated form of the full technical note (see *Flaaen* (2013c)) documenting the bridge between the DCA and the Business Register.

C.2.1.1 Directories of International Corporate Structure

The LexisNexis Directory of Corporate Affiliations (DCA) is the primary source of information on the ownership and locations of U.S. and foreign affiliates. The DCA describes the organization and hierarchy of public and private firms, and consists of three separate databases: U.S. Public Companies, U.S. Private Companies, and International – those parent companies with headquarters located outside the United States. The U.S. Public database contains all firms traded on the major U.S. exchanges, as well as major firms traded on smaller U.S. exchanges. To be included in the U.S. Private database, a firm must demonstrate revenues in excess of $1 million, 300 or more employees, or substantial assets. Those firms included in the International database, which include both public and private companies, generally have revenues greater than $10 million. Each database contains information on all parent company subsidiaries, regardless of the location of the subsidiary in relation to the parent.

The second source used to identify multinational firms comes from Uniworld Business Publications (UBP). This company has produced periodic volumes documenting the locations and international scope of i) American firms operating in foreign countries; and ii) foreign firms with operations in the United States. Although only published biennially, these directories benefit from a focus on multinational firms, and from no sales threshold for inclusion.

Because there exist no common identifiers between these directories and Census Bureau data infrastructure, we rely on probabilistic name and address matching — so-called "fuzzy merging" — to link the directories to the Census data infrastructure.

C.2.1.2 Background on Name and Address Matching

Matching two data records based on name and address information is necessarily an imperfect exercise. Issues such as abbreviations, misspellings, alternate spellings, and alternate name conventions rule out an exact merging procedure, leaving the researcher with probabilistic string matching algorithms that evaluate the "closeness" of match — given by a score or rank — between the two character strings in question. Due to the large computing requirements of these algorithms, it is common to use so-called "blocker" variables to restrict the search samples within each dataset. A "blocker" variable must match exactly, and as a result this implies the need for a high degree of conformity between these variables in the two datasets. In the context of name and address matching, the most common "blocker" variables are the state and city of the establishment.

The matching procedure uses a set of record linking utilities described in *Wasi and Flaaen* (2014). This program uses a bigram string comparator algorithm on multiple variables with differing user-specified weights.[1] This way the researcher can apply, for example, a larger weight on a near *name* match than on a perfect *zip code* match. Hence, the "match score" for this program can be interpreted as a weighted average of each variable's percentage of

[1]The term bigram refers to two consecutive characters within a string (the word *bigram* contains 5 possible bigrams: "bi", "ig", "gr", "ra", and "am"). The program is a modified version of *Blasnik* (2010), and assigns a score for each variable between the two datasets based on the percentage of matching bigrams. See *Flaaen* (2013c) or *Wasi and Flaaen* (2014) for more information.

bigram character matches.

C.2.1.3 The Unit of Matching

The primary unit of observation in the DCA, UBP, and BR datasets is the business establishment. Hence, the primary unit of matching is the establishment, and not the firm. However, there are a number of important challenges with an establishment-to-establishment link. First, the DCA (UBP) and BR may occasionally have differing definitions of the establishment. One dataset may separate out several operating groups within the same firm address (i.e. JP Morgan – Derivatives, and JP Morgan - Emerging Markets), while another may group these activities together by their common address. Second, the name associated with a particular establishment can at times reflect the subsidiary name, location, or activity (i.e. Alabama plant, processing division, etc), and at times reflect the parent company name. Recognizing these challenges, the primary goal of the matching will be to assign each DCA (UBP) establishment to the most appropriate business location of the parent firm identified in the BR. As such, the primary matching variables will be the establishment name, along with geographic indicators of street, city, zip code, and state.

C.2.1.4 The Matching Process: An Overview

The danger associated with probabilistic name and address procedures is the potential for false-positive matches. Thus, there is an inherent tension for the researcher between a broad search criteria that seeks to maximize the number of true matches and a narrow and exacting criteria that eliminates false-positive matches. The matching approach used here is conservative in the sense that the methodology will favor criteria that limit the potential for false positives at the potential expense of slightly higher match rates. As such, the procedure generally requires a match score exceeding 95 percent, except in those cases where ancillary evidence provides increased confidence in the match.[2]

This matching proceeds in an iterative fashion, in which a series of matching procedures are applied with decreasingly restrictive sets of matching requirements. In other words, the initial matching attempt uses the most stringent standards possible, after which the non-matching records proceed to a further matching iteration, often with less stringent standards. In each iteration, the matching records are assigned a flag that indicates the standard associated with the match.

See Table C.1 for a summary of the establishment-level match rate statistics by year and type of firm. Table C.2 lists the corresponding information for the Uniworld data.

C.2.1.5 Construction of Multinational Indicators

The DCA data allows for the construction of variables indicating the multinational status of the U.S.-based establishment. If the parent firm contains addresses outside of the United States, but is headquartered within the U.S., we designate this establishment as part of a U.S. multinational firm. If the parent firm is headquartered outside of the United States, we designate this establishment as part of a Foreign multinational firm. We also retain the nationality of parent firm.[3]

There can be a number of issues when translating the DCA-based indicators through the DCA-BR bridge for use within the Census Bureau data architecture. First, there may be disagreements between the DCA and Census on what constitutes a firm, such that an establishment matches may report differing multinational indicators for the same Census-identified firm. Second, such an issue might also arise due to joint-ventures. Finally, incorrect matches may also affect the degree to which establishment matches agree when aggregated to a firm definition. To address these issues, we apply the following rules when using the

[2] The primary sources of such ancillary evidence are clerical review of the matches, and additional parent identifier matching evidence.

[3] The multinational status of firms from the UBP directories are more straightforward.

DCA-based multinational indicators and aggregating to the (Census-based) firm level. There are three potential cases:[4]

Potential 1: A Census-identified firm in which two or more establishments match to different foreign-country parent firms

1. Collapse the Census-identified firm employment based on the establishment-parent firm link by country of foreign ownership
2. Calculate the firm employment share of each establishment match
3. If one particular link of country of foreign ownership yields an employment share above 0.75, apply that link to all establishments within the firm.
4. If one particular link of country of foreign ownership yields an employment share above 0.5 and total firm employment is below 10,000, then apply that link to all establishments within the firm.
5. All other cases require manual review.

Potential 2: A Census-identified firm in which one establishment is matched to a foreign-country parent firm, and another establishment is matched to a U.S. multinational firm.

1. Collapse the Census-identified firm employment based on the establishment-parent firm link by type of DCA link (Foreign vs U.S. Multinational)
2. Calculate the firm employment share of each establishment match
3. If one particular type of link yields an employment share above 0.75, apply that link to all establishments within the firm.

[4] Some of these cases also apply to the UBP-BR bridge.

4. If one particular type of link yields an employment share above 0.5 and total firm employment is below 10,000, then apply that link to all establishments within the firm.

5. All other cases require manual review.

Potential 3: A Census-identified firm in which one establishment is matched to a non-multinational firm, and another establishment is matched to a foreign-country parent firm (or U.S. multinational firm).

Apply same steps as in Potential 2.

C.2.2 Classifying Firm-Level Trade

The firm-level data on imports available in the LFTTD does not contain information on the intended use of the goods.[5] Disentangling whether an imported product is used as an intermediate input for further processing — rather than for final sale in the U.S. — has important implications for the nature of FDI, and the role of imported goods in the transmission of shocks. Fortunately, the Census Bureau data contains other information that can be used to distinguish intermediate input imports from final goods imports. Creating lists of the principal products produced by firms in a given detailed industry in the United States should indicate the types of products that, when imported, should be classified as a "final" good – that is, intended for final sale without further processing. The products imported outside of this set, then, would be classified as intermediate goods.[6] Such product-level production data exists as part of the "Products" trailer file of the Census of Manufacturers. As detailed in *Pierce and Schott* (2012) (see page 11), combining import, export, and

[5]This is one advantage of the survey data on multinational firms available from the Bureau of Economic Analysis. There are, however, a number of critical disadvantages of this data source, as outlined in *Flaaen* (2013a).

[6]To be more precise, this set will include a combination of intermediate and capital goods.

production information at a product-level is useful for just such a purpose.

C.2.2.1 Creating a NAICS-Based set of Final/Intermediate Products

As part of the quinquennial Census of Manufacturers (CM), the Census Bureau surveys establishments on their total shipments broken down into a set of NAICS-based (6 digit) product categories. Each establishment is given a form particular to its industry with a list of pre-specified products, with additional space to record other product shipments not included in the form. The resulting product trailer file to the CM allows the researcher to understand the principal products produced at each manufacturing establishment during a census year.

There are several data issues that must be addressed before using the CM-Products file to infer information about the relative value of product-level shipments by a particular firm. First, the trailer file contains product-codes that are used to "balance" the aggregated product-level value of shipments with the total value of shipments reported on the base CM survey form. We drop these product codes from the dataset. Second, there are often codes that do not correspond to any official 7-digit product code identified by Census. (These are typically products that are self-identified by the firm but do not match any of the pre-specified products identified for that industry by Census.) Rather than ignoring the value of shipments corresponding to these codes, we attempt to match at a more aggregated level. Specifically, we iteratively try to find a product code match at the 6, 5, and 4 digit product code level, and use the existing set of 7-digit matches as weights to allocate the product value among the 7-digit product codes encompassing the more aggregated level.

We now discuss how this file can be used to assemble a set of NAICS product codes that are the predominant output (final goods) for a given NAICS industry. Let x_{pij} denote the shipments of product p by establishment i in industry j during a census year. Then the total output of product p in industry j can be written as:

$$X_{pj} = \sum_{i=1}^{I_j} x_{pij},$$

where I_j is the number of firms in industry j. Total output of industry j is then:

$$X_j = \sum_{p=1}^{P_j} X_{pj}.$$

The share of industry output accounted for by a given product p is therefore:

$$S_{pj} = \frac{X_{pj}}{X_j}.$$

One might argue that the set of final goods products for a given industry should be defined as the set of products where $S_{pj} > 0$. That is, a product is designated as a "final good" for that industry if *any establishment* recorded positive shipments of the product. The obvious disadvantage of employing such a zero threshold is that small degrees of within-industry heterogeneity will have oversized effects on the classification.

Acknowledging this concern, we set an exogenous threshold level W such that any p in a given j with $S_{pj} > W$ is classified as a final good product for that industry. The upper portion of Table C.3 documents the number of final goods products and the share of intermediate input imports based on several candidate threshold levels. The issues of a zero threshold are quite clear in the table; a small but positive threshold value (0.1) will have a large effect on the number of products designated as final goods. This shows indirectly that there are a large number of products produced by establishments in a given industry, but a much smaller number that comprise the bulk of total value.

There are several advantages to using the CM-Products file rather than using an input-output table.[7] First, within a given CM year, the classification can be done at the firm or establishment level rather than aggregating to a particular industry. This reflects the fact that the same imported product may be used as an input by one firm and sold to consumers as a final product by another. Second, the CM-Products file is one of the principal data inputs into making the input-output tables, and thus represents more finely detailed information. Related to this point, the input-output tables are produced with a significant delay – the most recent available for the U.S. is for year 2002. Third, the input-output tables for the U.S. are based on BEA industry classifications, which imply an additional concordance (see below) to map into the NAICS-based industries present in the Census data.

We now turn to the procedure to map firm-level trade into intermediate and final goods using the industry-level product classifications calculated above.

C.2.2.2 Mapping HS Trade Transactions to the Product Classification

The LFTTD classifies products according to the U.S. Harmonized Codes (HS), which must be concorded to the NAICS-based product system in order to utilize the classification scheme from the CM-Products file. Thankfully, a recent concordance created by *Pierce and Schott* (2012) can be used to map the firm-HS codes present in the LFTTD data with the firm-NAICS product codes present in the CM-Products data.

A challenge of this strategy is that the LFTTD exists at a firm-level, while the most natural construction of the industry-level classification scheme is by establishment. More concretely, for multi-unit, multi-industry firms, the LFTTD is unable to decompose an import shipment into the precise establishment-industry of its U.S. destination. [8] While

[7]Another option is to use the CM-Materials file, the flip side of the CM-Products file. Unfortunately, the CM-Materials file contains significantly more problematic product codes than the Products file, and so concording to the trade data is considerably more difficult.

[8]It is worth pointing out that the most obvious way that this would materialize is by vertical integration of the firm in its U.S. operations. Provided that the industry designation of the firm pertains to its most downstream operations, then this is would not serve to bias the firms' classification of imported goods, as the upstream products are not actually "final" goods for that firm.

recognizing the caution that should be used in this regard, we adopt the approach that is commonly used in such circumstances: the industry of the firm is defined as that industry encompassing the largest employment share.

Once the firm-level trade data is in the same product classification as the industry-level filter created from the CM-Products file, all that is left is to match the trade data with the filter by NAICS industry. Thus, letting M_{ij} denote total imports from a firm i (firm i is classified as being in industry j), we can then categorize the firm's trade according to:

$$\left.\begin{aligned} M_{ij}^{\text{int}} &= \sum_{p \notin P_j} M_{ipj} \\ \\ M_{ij}^{\text{fin}} &= \sum_{p \in P_j} M_{ipj} \end{aligned}\right\} \qquad \text{where} \qquad P_j = \{p \mid S_{pj} \geq W\}\,. \tag{C.10}$$

The bottom section of Table C.3 shows some summary statistics of the intermediate share of trade according to this classification system, by several values of the product-threshold W. There are at least two important takeaways from these numbers. First, the share of intermediates in total imports is roughly what is reported in the literature using IO Tables. Second, the share of total trade occupied by intermediate products is not particularly sensitive to the exogenous threshold level. While there is a small increase in the share when raising the threshold from 0 to 0.1 (about 3 percentage points), the number is essentially unchanged when raising it further to 0.2.

C.2.3 Sample Selection

C.2.3.1 Constructing the Baseline Dataset

This section will discuss the steps taken to construct the sample used in section 3.3.1.

Beginning with the raw files of the LFTTD export/import data, we drop any transactions with missing firm identifiers, and those pertaining to trade with U.S. territories. Next, we merge the LFTTD files with the HS-NAICS6 product concordance from *Pierce and Schott* (2012); if there is no corresponding NAICS6 code for a particular HS code, then we set NAICS6 equal to XXXXXX. We then aggregate up to the level of Firm-Country-Month-NAICS6, and then create extracts according to three sets of destinations/sources: Japan, Non-Japan, and North America (Canada and Mexico). Then, assigning each firm to an LBD-based industry (see below), we run the NAICS-based trade codes through the intermediate/final goods filter discussed in Appendix C.2.2. The firms' monthly trade can then be split into intermediate and final goods components. We repeat this step for years 2009, 2010, and 2011.

Using the Longitudinal Business Database, we drop inactive, ghost/deleted establishments, and establishments that are not in-scope for the Economic Census. To create the sample of manufacturing firms in the U.S., we first create a firm industry code defined as the industry encompassing the largest share of firm employment. We then drop non-manufacturing firms. Next, we merge the LBD for each year with the DCA-Bridge (see section C.2.1) containing multinational indicators. We then apply the rules specified above for clarifying disagreements with the DCA-based multinational indicators. After creating monthly copies of each firm, we merge by firm-month to the trade data. Missing information of trade data is altered to represent zeros. We repeat these steps for years 2009-2011, and then append the files together. Firms that do not exist in all three years are dropped from the sample.

C.2.3.2 GIS Mapping of Earthquake Intensity Measures to Affiliate Locations

As part of the Earthquake Hazards Program, the U.S. Geological Survey produces data and map products of the ground motion and shaking intensity following major earthquakes. The preferred measure to reflect the perceived shaking and damage distribution is the estimated "Modified Mercalli Intensity (MMI)" which is based on a relation of survey response and measured peak acceleration and velocity amplitudes. The USGS extends the raw data from geologic measurement stations and predicts values on a much finer grid using standard seismological inferences and interpolation methods. The result is a dense grid of MMI values covering the broad region affected by the seismic event. For more information on this methodology, see *Wald et al.* (2006).

To utilize this information, we take all Japanese addresses from the DCA/Uniworld directories that correspond to any U.S. operation via an ownership link. We geocode these addresses into latitude/longitude coordinates using the Google Geocoding API, and then compute the inverse distance-weighted mean of the relevant seismic intensity measure based on a 10km radius surrounding a given establishment. The firm identifiers within the corporate directories allow us to create firm-specific measures (average and maximum values, by manufacturing/non-manufacturing), which can then be brought into the baseline Census dataset via the bridges discussed in appendix C.2.1.

Table C.1: DCA Match Statistics: 2007-2011

	# of DCA Establishments	Matched to B.R.	Percent Matched
Total			
2007	112,346	81,656	0.73
2008	111,935	81,535	0.73
2009	111,953	81,112	0.72
2010	111,998	79,661	0.71
2011	113,334	79,516	0.70
U.S. Multinationals			
2007	22,500	16,396	0.73
2008	23,090	16,910	0.73
2009	22,076	16,085	0.73
2010	21,667	15,785	0.73
2011	21,721	15,557	0.72
Foreign Multinationals			
2007	10,331	7,555	0.73
2008	9,351	6,880	0.74
2009	11,142	8,193	0.74
2010	11,308	8,181	0.72
2011	11,619	8,357	0.72

Table C.2: Uniworld Match Statistics: 2006-2011

	# of Uniworld Establishments	Matched to B.R.	Percent Matched
Foreign Multinationals			
2006	3,495	2,590	0.74
2008	3,683	2,818	0.76
2011	6,188	4,017	0.65
U.S. Multinationals[1]			
2007	4,043	3,236	0.80
2009	4,293	3,422	0.80

[1]U.S. multinationals include only the establishment identified as the U.S. headquarters.

Table C.3: Appendix Table Comparing the Results from Threshold Values W

	Threshold Values		
	$W = 0$	$W = 0.1$	$W = 0.2$
Number of Final Good Products per Industry			
Median	19	1	1
Mean	25	1.52	1.14
Min	1	1	0
Max	154	6	3
Implied Share of Intermediate Inputs			
Imports	60.9	63.90	63.97
Exports	52.0	54.96	55.04

C.3 Appendix: Other Results

C.3.1 Alternate Specifications for Treatment Effects Regressions

Our results from section 3.3.2 are based on a sample including all Japanese multinationals in manufacturing, and therefore uses a levels specification to allow for zeros in the firm-month observations. Because larger firms exhibit greater absolute deviations from trend, this roughly amounts to weighting firms based on size, such that the results correspond to a representative firm based on the aggregate effect of the group.

To see this, and to explore how the levels specification influences our interpretation, we repeat the analysis on a subset of the firms for which we can view the percentage changes directly. Specifically, we drop any firms with zeros in any month for intermediate imports or N.A. exports during the sample, and then take logs and HP-filter each series to obtain percentage deviations from trend for each firm.[9] The results of this exercise are shown in Panel A of Figure C.1. We suppress standard errors for the sake of clarity; the drops are significant at the 95% level for between 2-4 months following the shock. If we rerun these regressions while also weighting according to the pre-shock size of firms, we obtain a picture that looks much closer to Figure 3.7, see Panel B of Figure C.1.

These results indicate that the larger firms appear to be affected the most from this shock. This could be partly a result of our proxy being less effective for smaller firms that may not engage in consistent exports to North America.

C.3.2 Probit Model of Import/Output Disruptions

We specify a simple probit model to understand the relative importance of various firm-level characteristics in the import and output declines following the tsunami. The model is

[9]We re-weight the control group as described in section 3.3.1.

$$Pr(X_{ik}^D = 1) = \Phi\left[\beta_1 \text{JPN}_{ik} + \beta_2 Exposed_{ik} + \beta_3 MMI_{ik} + \beta_4 Port_{ik} + \gamma_k\right] \qquad \text{(C.11)}$$

where the dependent variable (X_{ik}^D) is an indicator equal to one if the N.A. exports of firm i in industry k are on average 20% below trend during the five months following the Tōhoku event. The independent variables are also indicators: JPN_{ik}, for affiliates of Japanese multinationals; $Exposed_{ik}$, for firms with an exposure to Japanese inputs above 0.05 of total material; MMI_{ik} for firms with an elevated MMI value pertaining to their average Japanese manufacturing locations; and $Port_{ik}$ for firms that typically rely on imports via ports damaged by the tsunami.[10] The γ_k term allows for industry-specific intercepts. To evaluate the determinants of an input disruption from Japan, we replace the dependent variable with J_{ik}^D, an indicator for a drop in Japanese imported inputs of 20% relative to trend.

Panel A of Table C.4 evaluates firm characteristics predicting a drop in U.S. output (X_{ik}^D), as measured by our proxy. The columns (1)-(4) show the results from different specifications with various combinations of the covariates in equation (C.11). Both Japanese ownership and high exposure to Japanese inputs significantly increase the probability of an output disruption, as expected. In columns (3) and (4), we demonstrate that Japanese ownership is substantially more indicative of an output decline than high input exposure alone. In Panel B, we replace the dependent variable with the binary measure of a drop in Japanese intermediate inputs (J_i^D). The results from these regressions indicate, unsurprisingly, that high exposure to Japanese imports are highly predictive of a subsequent disruption following the Tōhoku event. Apart from their exposure to imports from Japan, the Japanese affiliates

[10]Specifically, the $MMI_{ik} = 1$ if the average Japanese manufacturing establishment corresponding to a U.S. firm is above the median (roughly an MMI of 5.2) of all firms with Japanese manufacturing locations. The affected ports are: Onahama, Hitachi, Kashima, Haramachi, Shiogama, Sendai, Shimizu, Ishinomaki, Hashinohe, Miya Ko, Kamaishi, Ofunato, and Kessennuma.

are no more likely to suffer a disruption to these imports (see column 8).[11] While the results from Table C.4 are somewhat inconclusive, they nevertheless point to unique features of the production function of Japanese affiliates that yields direct pass-through of Japanese shocks to the U.S. economy. Our estimation procedure that follows should help to clarify this point further.

C.3.3 Bootstrapping Standard Errors

We use bootstrapping methods to compute measures of the dispersion of our point estimates. Using random sampling with replacement within each group of firms, we create 5000 new artificial samples and re-run the estimation procedure. The standard deviation of the point estimates across these bootstrap samples is shown in Table 3.3. To gain a more complete picture of the dispersion, we create density estimates for each sample of firms across the parameter space for the elasticities. These densities are shown in Figure C.3.

C.3.4 Effects on U.S. Exports to Japan

Another dimension of the transmission of the Tōhoku shock to the United States is U.S. exports back to Japan. To the extent that firms in the U.S. receive inputs from Japan for processing and re-shipment back to Japan, one might expect the U.S. exports to Japan may fall following the Tōhoku event. On the other hand, U.S. firms may have increased shipments to Japan following the shock in order to offsett what were large production and supply shortages within Japan. To evaluate this, we re-run the specification in equation (3.5) but replace $V^M_{i,t}$, the value of intermediate imports of firm i in month t, with $V^{JEXP}_{i,t}$, the value of Japanese exports of firm i in month t. The results are shown in Figure C.2. As is clear from the figure, we do not see strong evidence to support either hypothesis regarding this particular trade flow, at least as it pertains to Japanese multinationals in particular.

[11] The combined effect of the coefficients on Japan and JPN*Exp is -0.16, and not significant.

C.3.5 Effects on Employment and Payroll

The Standard Statistical Establishment List (SSEL) contains quarterly employment and payroll information for all employers (with some small exceptions) in the U.S. economy. This list is held separately as a single-unit(SSEL-SU) and multi-unit (SSEL-MU) file. The Report of Organization Survey (ROS) asks firms to list the establishments which report under a particular EIN, and this information is then recorded to the firm identifier on the Multi-Unit File. To build a quarterly employment series at the firm-level, we link the EIN variables on the SU file with the firm-identifier linked with each EIN on the MU file. In principle, the four quarters of payroll listed on the SSEL is combined by Census to create an annual payroll figure for each establishment, which is the value recorded in the LBD. Similarly, the employment variable corresponding to the 1st quarter (week of March 12) from the SSEL is that used by the LBD.

Once we merge the SSEL-based data with quarterly employment and payroll to the LBD for a particular year, we conduct a series of reviews to ensure that the annual payroll (and 1st quarter employment) roughly align. Any establishments with disagreements between the SSEL-based payroll and LBD-based payroll such that the ratio was greater than 2 or less than 0.6 were dropped.

After these modifications were made, the remainder of the data construction was similar to that in section C.2.3. We merge multinational indicators from the DCA, drop non-manufacturing firms, append the 2009, 2010, and 2011 files together, and keep only those firms that exist in each year. Using the same set of firms as a control group as specified in section 3.3.1, we run the following regression:

$$\Delta \text{emp}_{j,t} = \sum_{i=-3}^{3} \gamma_i E_i + \sum_{i=-3}^{3} \beta_i E_i D_{j,i} + u_{j,t} \tag{C.12}$$

where $\Delta \text{emp}_{j,t} \equiv ln(\text{emp}_{j,t}/\text{emp}_{j,t-4})$, where $\text{emp}_{j,t}$ indicates employment at firm j in quarter t. We also re-run the equation specified in equation C.12 using payroll $pay_{j,t}$ as the dependent variable (where $\Delta \text{pay}_{j,t} \equiv ln(\text{pay}_{j,t}/\text{pay}_{j,t-4})$. The qualitative results are shown in table C.5.

C.3.6 Effects on Unit Values (Prices) of Trade

The LFTTD contains information on quantities as well as values for each trade transaction, recorded at a highly disaggregated product definition (HS 10 digit). This allows for the construction of unit values (prices) for each firm-product-month observation, which allows for an analysis of price movements surrounding the Tōhoku event.

The majority of the data construction is identical to that in section C.2.3, however there are a number of modifications. First, we drop all transactions with missing or imputed quantities in the LFTTD, and then aggregate to the Firm-HS10-month frequency, separately for each type of trade transaction: 1) Related-Party imports from Japan; 2) Non Related-Party imports from Japan; 3) Related-Party exports to Canada/Mexico; and 4) Non Related-Party exports to Canada/Mexico. Next, we select only those firms identified as manufacturing in the LBD. We keep the related-party and arms-length transactions separate as one may expect these prices to behave differently following a shock. As above, we keep only manufacturing firms, append the annual files together, and then select only those firms identified as a multinational in either 2009, 2010, or 2011.

At the product level, there is little reason to suspect trends or seasonal variation over this short of a time period. Moreover, there is no concern here about accounting for zeros in the data. As such we take a firm j's imports (exports) of product p in month t, and run the following specification in logs ($\text{m}_{p,j,t} = log(\text{M}_{p,j,t})$:

$$\text{m}_{p,j,t} = \alpha_{pj} + \sum_{i=-19}^{9} \gamma_i E_i + \sum_{i=-19}^{9} \beta_i E_i D_{j,i} + u_{j,t} \tag{C.13}$$

where α_{pj} are firmXproduct fixed-effects, γ_i are monthly fixed effects (with the dummy variable $E_i's$ corresponding to each calendar month), and $u_{j,t}$ are random effects. The variables $D_{j,t}$ are dummy variables equal to one if the firm is owned by a Japanese parent company.

A qualitative version of the results is shown in Table C.6.

C.3.7 Ward's Automotive Data

Ward's electronic databank offers a variety of data products for the global automotive industry at a monthly frequency. We obtain Japanese production (by model), North American production (by plant and model), U.S. inventory (by model), and North American sales (by model) all for the period January 2000 to December 2012. The inventory and sales data also contain the country of origin, so one can separate out these variables based on whether a particular model was imported vs domestically-produced. The series cover the universe of the assembly operations of finished cars and light trucks. Unfortunately, there is no information on input shipments.

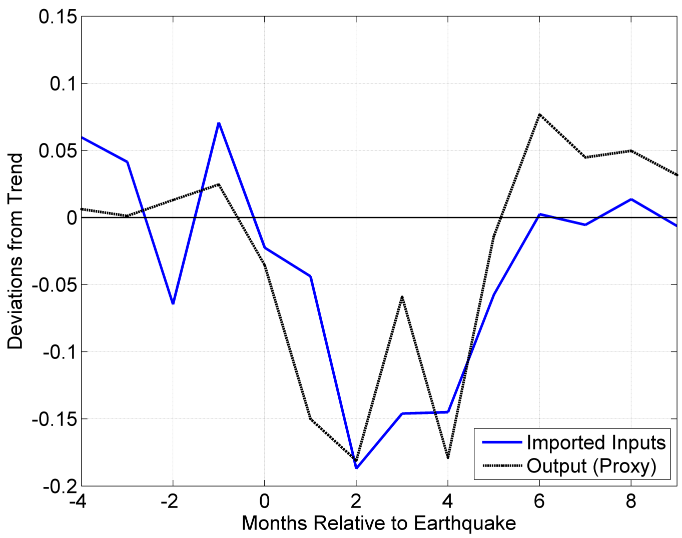
For the plant-level analysis of production, the base sample consists of 167 plants active at some point during 2000-2012. We remove plants that were not continuously in operation during the period 2009-2012, and combine several plants that are recorded separately in the data, but are in effect the same plant. After these modifications, the sample reduces to 62 plants, 22 of which are owned by a Japanese parent. The average monthly production in the three months preceding the shock is 12,904 for Japanese plants, and 14,903 for Non-Japanese plants. The specification is identical to that in section 3.3.1:

$$Q_{i,t} = \alpha_0 + \alpha_i + \sum_{p=-14}^{9} \gamma_p E_p + \sum_{p=-14}^{9} \beta_p E_p \mathrm{JPN}_{i,p} + u_{i,t} \tag{C.14}$$

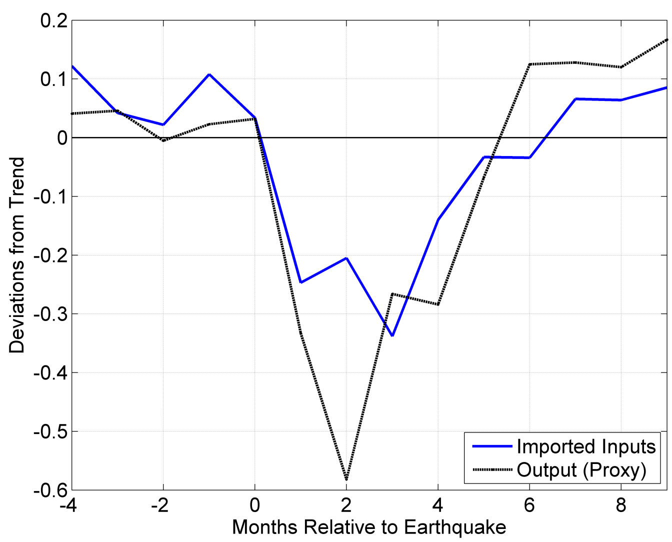
where here the variable $Q_{i,t}$ is auto production by plant i in month t, after removing a plant-specific trend though March 2011. Because these plants can be tracked with some confidence back in time, it is reasonable here to remove seasonality directly, rather than assume a shared seasonal component between the treated and control groups as in section 3.3.2. We use the X12-ARIMA model, provided by the National Bank of Belgium, and apply it to each series before correcting for trend. The results for the Japanese plants are mostly similar, as shown in table C.7.

Figure C.1: Relative Inputs and Output (Proxy) of Japanese Firms (Reduced Sample) Logged, HP-Filtered

A. No Size-Weighting

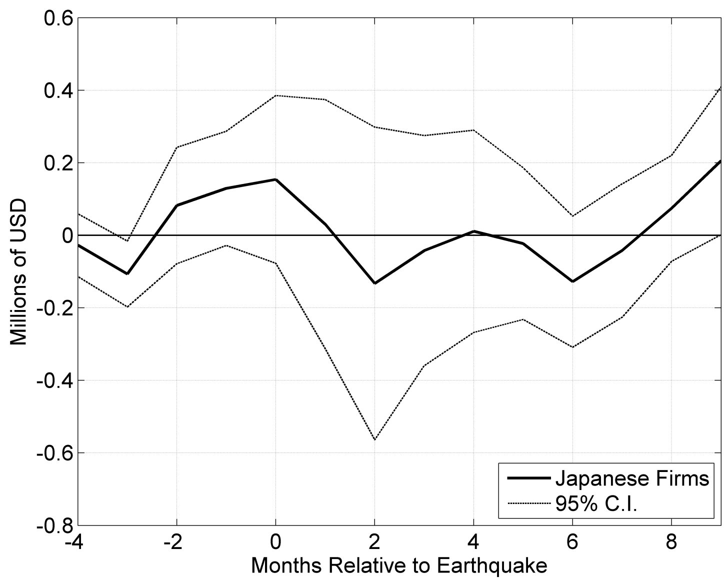


B. Size-Weighted



Source: LFTTD-DCA-UBP as explained in text.
These figures report the relative percentage deviations from trend of Japanese affiliates relative to a control group of other multinational firms. The values are coefficient estimates taken from an interaction of a Japanese-firm dummy with a monthly dummy – additional baseline monthly dummies remove seasonal effects. These results reflect a reduced sample with no firm-month zeros in imported inputs or N.A. exports. The data is logged, and HP-filtered using a monthly smoothing parameter.

Figure C.2: Dynamic Treatment Effects: Relative Japanese Exports of Japanese Firms



Source: LFTTD-DCA-UBP as explained in text.
These figures report the Japanese exports of the U.S. affiliates of Japanese firms relative to a control group of other multinational firms. The values are coefficient estimates taken from an interaction of a Japanese-firm dummy with a monthly dummy – additional baseline monthly dummies remove seasonal effects. Standard errors are clustered at the firm level.

Table C.4: Predicting Japanese Import and U.S. Output Disruption by Firm Characteristics

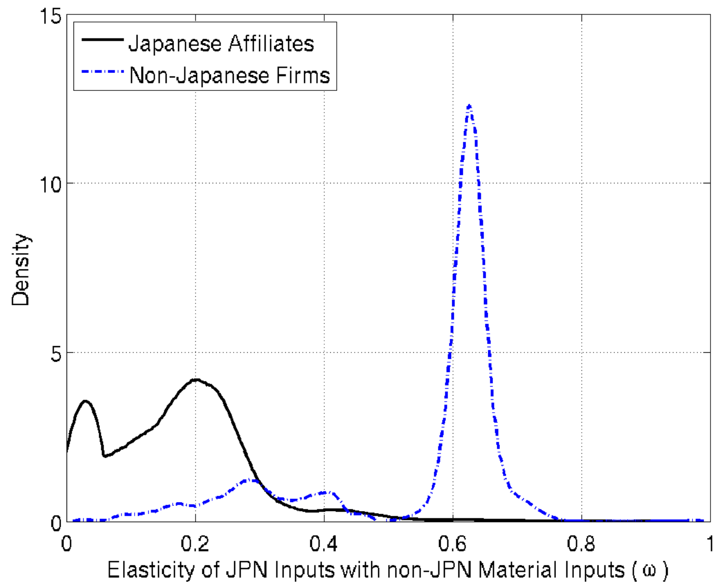
	Panel A: Disruption to U.S. Output (proxy)				**Panel B: Disruption to Japanese Imports**			
	$X_i^D = 1$				$J_i^D = 1$			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Japan	0.443***		0.352***	0.347**	0.707***		0.310***	0.686***
	(0.0921)		(0.117)	(0.152)	(0.0917)		(0.115)	(0.150)
Exposed		0.351***	0.145	0.140		0.814***	0.636***	0.991***
		(0.0886)	(0.112)	(0.149)		(0.0880)	(0.110)	(0.144)
JPN*Exp				-0.00771				-0.848***
				(0.228)				(0.222)
MMI	-0.176***	-0.121*	-0.178***	-0.178***	0.346***	0.389***	0.341***	0.306***
	(0.0676)	(0.0646)	(0.0676)	(0.0683)	(0.0691)	(0.0667)	(0.0694)	(0.0704)
Ports	-0.174	-0.144	-0.197		0.248	0.217	0.168	0.174
	(0.224)	(0.225)	(0.226)		(0.211)	(0.212)	(0.213)	(0.213)
Constant	-0.674	-0.674	-0.674	-0.674	-4.672	-4.672	-4.672	-4.668
	(0.681)	(0.681)	(0.681)	(0.681)	(85.78)	(85.78)	(85.78)	(85.00)
Industry Dummies	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations	2451	2451	2451	2451	2451	2451	2451	2451

*** p< 0.01, **p< 0.05, * p< 0.1

Source: LFTTD, DCA, UBP, and USGS as explained in the text. This table reports the results of a probit model prediction of JPN import and N.A. exports (output) disruption based on firm characteristics. See section 3.3.1 for a definition of the variables.

Figure C.3: Density Estimates of Elasticities Across Bootstrap Samples

A. Japanese vs non-Japanese Multinationals: Materials Elasticity (ω)



B. Japanese vs non-Japanese Multinationals: Materials-Capital/Labor Elasticity (ζ)

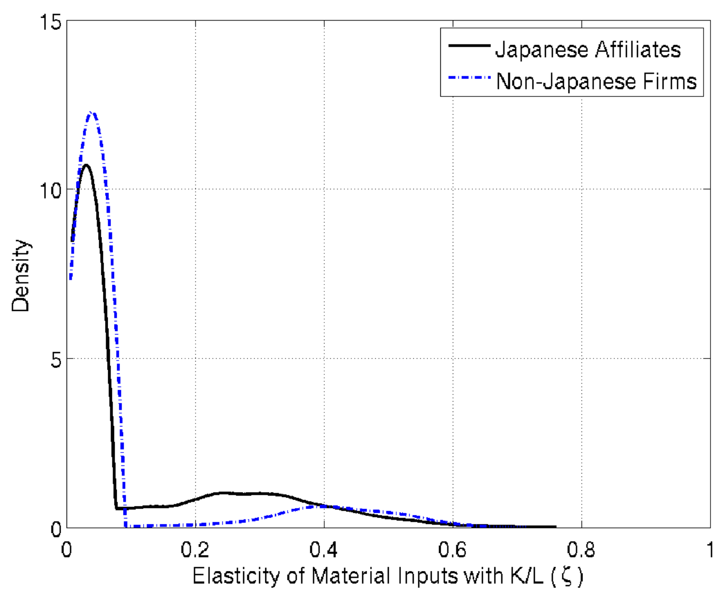
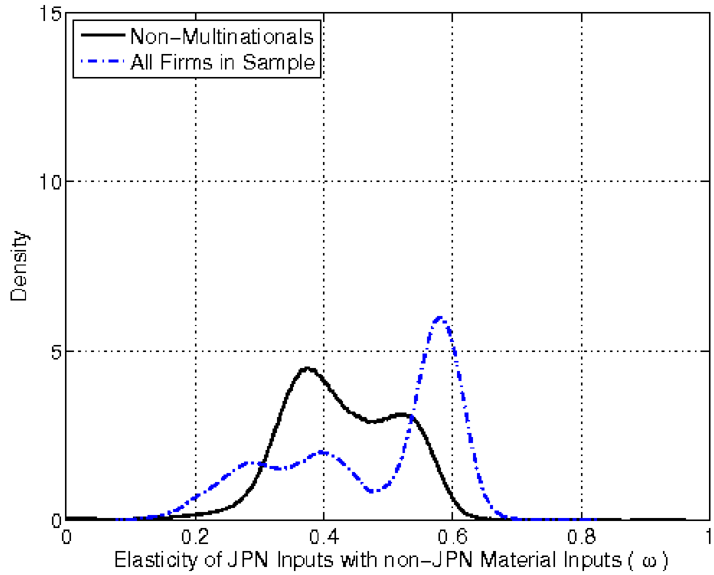
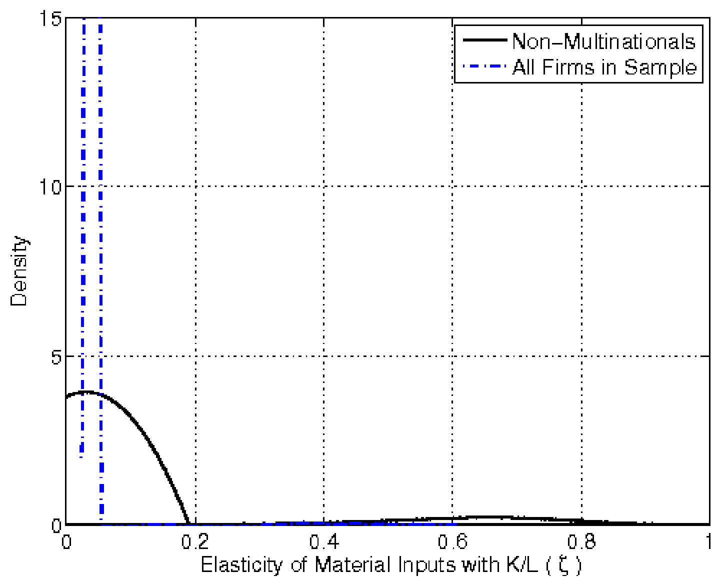


Figure C.3: Density Estimates of Elasticities Across Bootstrap Samples

C. Non-multinationals and All Firms: Materials Elasticity (ω)



D. Non-multinationals and All Firms: Materials-Capital/Labor Elasticity (ζ)



Source: LFTTD-DCA-UBP as explained in text.

Table C.5: Dynamic Treatment Effects: Quarterly Employment/Payroll Surrounding Tōhoku Event

	Log 4-Quarter Difference	
	Employment	Payroll
Independent Variables	(1)	(2)
Q2_2010 (t=-3)	pos***	pos***
Q3_2010 (t=-2)	pos***	pos***
Q4_2010 (t=-1)	pos***	pos***
Q1_2011 (t=0)	pos***	pos***
Q2_2011 (t=1)	pos***	pos***
Q3_2011 (t=2)	pos***	pos***
Q4_2011 (t=3)	pos***	pos***
JPNxQ2_2010 (t=-3)	neg	neg
JPNxQ3_2010 (t=-2)	neg	neg
JPNxQ4_2010 (t=-1)	neg	neg
JPNxQ1_2011 (t=0)	neg	neg
JPNxQ2_2011 (t=1)	neg	neg
JPNxQ3_2011 (t=2)	neg	neg
JPNxQ4_2011 (t=3)	neg	pos
constant	neg***	neg***
Firm Fixed Effects	Yes	Yes
Observations		
R-squared		

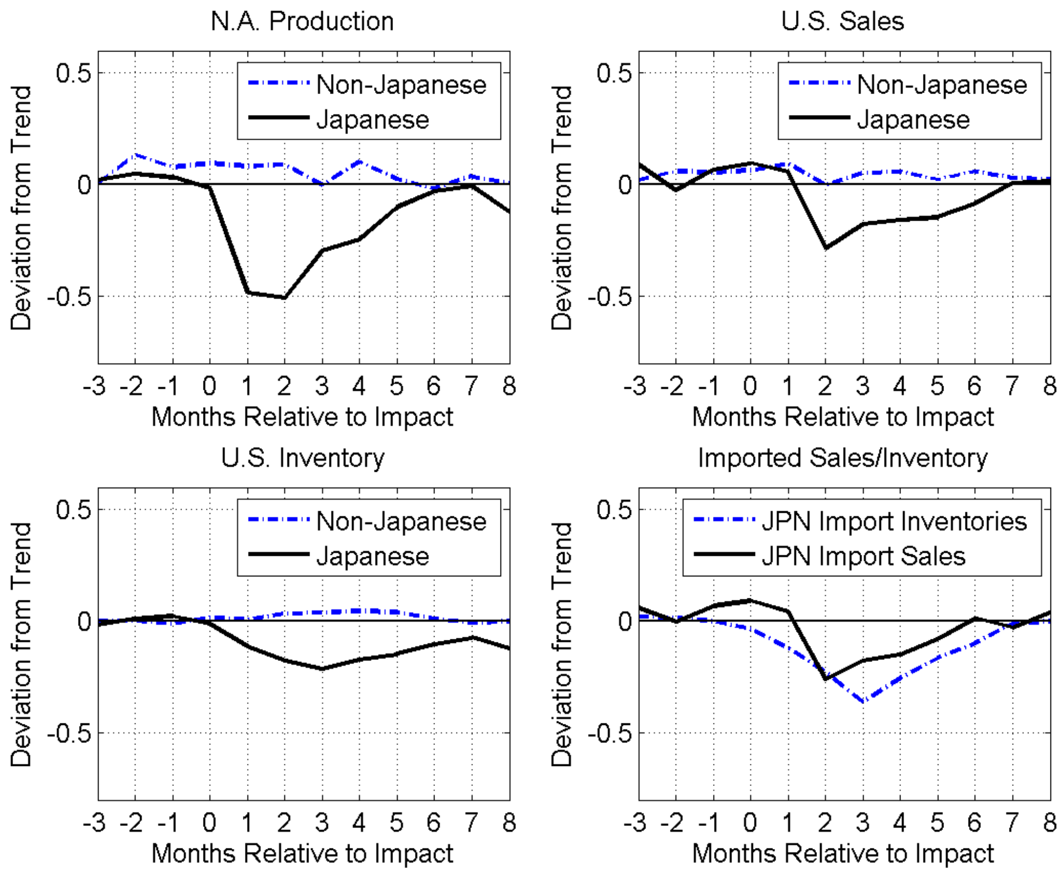
Source: SSEL and DCA as explained in the text.
Robust standard errors (clustered at the firmXProduct level) pertaining to each sign coefficient are indicated by: *** p<0.01, ** p<0.05, * p<0.1.
This table reports qualitative features of firm employment and firm payroll in the quarters surrounding the Tōhoku earthquake and tsunami. The first set of coefficients correspond to quarter dummies, whereas the second set (JPNx) correspond to the interaction of a Japanese firm dummy with quarter dummies. See equation C.12 in the text. The dependent variable is the four-quarter log difference of employment (payroll).

Table C.6: Dynamic Treatment Effects: Unit Values of Trade Surrounding Tōhoku Event

	Log Unit-Value of:			
	JPN Imports: Related Party	JPN Imports: Non-Related Party	N.A. Exports Related Party	N.A. Exports Non-Related Party
Independent Variables	(1)	(2)	(3)	(4)
Sep 2010 (t=-6)	neg**	pos	pos*	pos
Oct 2010 (t=-5)	pos	neg	pos**	pos
Nov 2010 (t=-4)	pos	pos	pos**	pos
Dec 2010 (t=-3)	pos	neg	pos	pos
Jan 2011 (t=-2)	neg	pos	neg	pos
Feb 2011 (t=-1)	pos	neg	pos**	pos
Mar 2011 (t=0)	neg	pos	pos	pos
Apr 2011 (t=1)	pos	pos	pos	pos
May 2011 (t=2)	neg	pos	neg	pos**
Jun 2011 (t=3)	pos**	neg	pos**	neg
Jul 2011 (t=4)	neg	neg	pos	neg
Aug 2011 (t=5)	pos	pos	neg	pos
Sep 2011 (t=6)	pos	pos	pos	pos**
Oct 2011 (t=7)	neg	neg	pos	pos
Nov 2011 (t=8)	pos	neg	pos	neg
Dec 2011 (t=9)	neg	pos	pos**	pos
JPNxSep 2010 (t=-6)	pos**	neg*	neg**	neg
JPNxOct 2010 (t=-5)	neg*	pos	pos	pos
JPNxNov 2010 (t=-4)	neg	pos	neg	neg
JPNxDec 2010 (t=-3)	neg	neg*	pos	pos
JPNxJan 2011 (t=-2)	pos	neg	neg	neg
JPNxFeb 2011 (t=-1)	neg	pos	pos	pos**
JPNxMar 2011 (t=0)	pos	pos	neg	neg
JPNxApr 2011 (t=1)	neg	pos	neg	neg
JPNxMay 2011 (t=2)	pos	neg	pos	neg
JPNxJun 2011 (t=3)	neg	pos*	neg	neg
JPNxJul 2011 (t=4)	pos	neg	pos	neg
JPNxAug 2011 (t=5)	neg*	neg*	neg	pos
JPNxSep 2011 (t=6)	neg	neg	neg	neg
JPNxOct 2011 (t=7)	pos	neg	neg	neg
JPNxNov 2011 (t=8)	neg	neg	neg	pos
JPNxDec 2011 (t=9)	neg	neg	pos	neg
constant	pos	neg	neg	neg
FirmXProduct Fixed Effect	Yes	Yes	Yes	Yes
Observations				
R-Squared				

Source: LFTTD, DCA, and UBP as explained in the text.
Robust standard errors (clustered at the firmXProduct level) pertaining to each sign coefficient are indicated by: *** p<0.01, ** p<0.05, * p<0.1 .
This table reports qualitative features of the unit values of trade surrounding the 2011 Tōhoku earthquake and tsunami. The first set of coefficients correspond to monthly dummies, whereas the second set (JPNx) correspond to the interaction of a Japanese firm dummy with monthly dummies. See equation C.13 in the text.

Figure C.4: Automotive Production, Inventory, Sales by Firm Type, Distributed Lag Model



Source: Ward's Automotive Database
This figure reports North American production, and U.S. sales and inventory data according to firm type: Japanese and non-Japanese firms. The values are coefficient estimates taken from a distributed lag model, exploiting time-series variation only. The underlying series have been seasonally adjusted, logged, and HP-Filtered Standard errors are suppressed in the interests of clarity. The Japanese automakers are Honda, Mazda, Mitsubishi, Nissan, Toyota, and Subaru.

Table C.7: Dynamic Treatment Effects: N.A. Automotive Production

VARIABLES	(1) Prod	(2) Prod	VARIABLES (cont'd)	(1) Prod (cont'd)	(2) Prod (cont'd)
Nov 2010 (t=-4)	91.06	17.78	JPN x Nov 2010 (t=-4)	-195.8	-341.7
	(649.9)	(608.8)		(841.9)	(799.2)
Dec 2010 (t=-3)	-1,973***	310.3	JPN x Dec 2010 (t=-3)	-385.0	-408.3
	(467.5)	(497.5)		(736.5)	(706.4)
Jan 2011 (t=-2)	-611.5	1,083*	JPN x Jan 2011 (t=-2)	781.0	-1,092
	(637.3)	(618.7)		(792.1)	(804.6)
Feb 2011 (t=-1)	694.9*	756.3*	JPN x Feb 2011 (t=-1)	-1,142	-1,210*
	(401.9)	(394.7)		(696.2)	(666.8)
Mar 2011 (t=0)	4,356***	1,483***	JPN x Mar 2011 (t=0)	-3,515***	-2,592***
	(524.9)	(389.1)		(812.0)	(842.7)
Apr 2011 (t=1)	-216.2	305.5	JPN x Apr 2011 (t=1)	-6,239***	-6,099***
	(707.7)	(620.4)		(1,303)	(1,282)
May 2011 (t=2)	1,584***	799.1	JPN x May 2011 (t=2)	-7,244***	-6,625***
	(525.4)	(511.3)		(1,651)	(1,740)
Jun 2011 (t=3)	1,366**	-499.3	JPN x Jun 2011 (t=3)	-4,564***	-3,423**
	(623.6)	(594.9)		(1,248)	(1,320)
Jul 2011 (t=4)	-4,512***	123.3	JPN x Jul 2011 (t=4)	-2,143	-3,723***
	(878.4)	(606.2)		(1,430)	(1,045)
Aug 2011 (t=5)	685.6	-1,323**	JPN x Aug 2011 (t=5)	-1,275	-1,108
	(744.0)	(648.1)		(970.8)	(1,012)
Sep 2011 (t=6)	-836.5	-1,895***	JPN x Sep 2011 (t=6)	-359.4	40.37
	(663.7)	(641.5)		(930.7)	(959.8)
Oct 2011 (t=7)	-338.0	-1,434**	JPN x Oct 2011 (t=7)	93.27	-265.4
	(662.3)	(632.4)		(885.6)	(785.8)
Nov 2011 (t=8)	-1,393**	-1,443**	JPN x Nov 2011 (t=8)	-1,318	-2,059*
	(582.8)	(601.2)		(1,159)	(1,183)
Dec 2011 (t=9)	-4,511***	-1,619**	JPN x Dec 2011 (t=9)	759.1	24.95
	(774.4)	(655.5)		(1,105)	(803.9)
Constant				-1,535***	-1,683***
				(89.30)	(91.95)
Plant Fixed Effects				Yes	Yes
Remove Plant-Specific Pre-Shock Trend				Yes	Yes
Remove Seasonal Component				No	Yes
Observations				2,976	2,976
R-squared				0.260	0.272

Source: Ward's Automotive Yearbook

Robust standard errors (clustered at the plant level) in parentheses. *** p<0.01, ** p<0.05, * p<0.1 .

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