

MICRO MPCs AND MACRO COUNTERFACTUALS:  
THE CASE OF THE 2008 REBATES\*

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We present evidence that the high estimated MPCs from the leading household studies result in implausible macroeconomic counterfactuals. Using the 2008 tax rebate as a case study, we calibrate a standard macro model with the estimated micro MPCs to construct counterfactual macroeconomic consumption paths in the absence of a rebate. The counterfactual paths imply that consumption expenditures would have plummeted in spring and summer 2008 and then mostly recovered in September 2008. We use narratives and forecasts to argue that these paths are implausible. We then show that standard two-way fixed effect estimates of the micro MPCs are upward biased. When we correct for the biases, we estimate smaller micro MPCs using the CEX data than the previous literature. We also show that realistic modifications of the model result in general equilibrium forces that dampen rather than amplify micro MPCs. The combination of smaller micro MPCs and dampening general equilibrium forces implies general equilibrium consumption multipliers that are below 0.2. JEL codes: E21, E27, E62.

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# I Introduction

Numerous studies have used household data to estimate the marginal propensity to consume out of temporary changes in income. Some of the leading studies in this area estimate the effects of the temporary U.S. tax rebates of 2001 and 2008. For example, Shapiro and Slemrod (2003, 2009), Johnson, Parker, and Souleles (2006), Sahm, Shapiro, and Slemrod (2010, 2012), Parker, Souleles, Johnson, and McClelland (2013), and Broda and Parker (2014) are exemplars in entrepreneurial data collection and the use of natural experiments to obtain estimates of this key micro parameter of interest to macroeconomists. Parker and co-authors found some very high estimates for the marginal propensity to consume (MPC). For example, Parker, Souleles, Johnson, and McClelland (2013) found a marginal propensity to spend out of the temporary tax rebate of 50 to 90 percent on total consumption within three months of receiving the 2008 tax rebate (p. 2531 and Table 3).

Estimates from these and other household-level studies have motivated the literature on heterogeneous agent models in which some households live hand to mouth because of myopia or financial market imperfections. The estimates have been used to calibrate a wide variety of macro New Keynesian heterogeneous agent models and to argue that temporary tax rebates can have large aggregate multipliers. For example, Kaplan and Violante (2014), Huntley and Michelangeli (2014), and Kaplan, Moll, and Violante (2018) calibrate their models to match a quarterly MPC of 0.25 on the nondurables component of consumption expenditures whereas Auclert, Rognlie, and Straub (2024) calibrate to an annual MPC of 0.5 and higher. Government policy in recent years has been guided by the high MPC estimates. For example, the Federal Reserve Board Staff assumed a six-month MPC of 0.5 in their March 2008 Greenbook (Edelberg and Sahm, 2008).

In this paper, we present evidence that the high estimated MPCs from the leading household studies result in implausible macroeconomic counterfactuals. Using the 2008 tax rebate as a case study, we calibrate a standard medium-scale New Keynesian model with the estimated MPCs to construct counterfactual macroeconomic consumption paths had there been no rebate. The paths imply a deep V-shape in aggregate consumption expenditures in spring and summer 2008. Based on narrative evidence and forecasts, we argue that these paths are implausible and that the actual stimulus effect of the rebates must have been modest. In the second half of the paper, we rec-

oncile the micro estimates with the aggregate data by correcting for upward biases in the micro MPC estimates and by introducing realistic dampening forces in the macro model.

More generally, our analysis of the 2008 rebate illustrates the use of historical macro counterfactuals as a diagnostic tool to assess the plausibility of micro estimates and macro models. In Orchard, Ramey, and Wieland (Forthcoming), we advocate the wider use of plausibility analysis and apply our methods to the 2001 rebate. Ramey (2025) applies the method to episodes in two small open economies, Singapore and Australia.

Some earlier work did question the high MPC estimates from the 2008 rebate based on aggregate data. In their analyses of the aggregate effects of the tax rebates of 2008, Feldstein (2008) and Taylor (2009) found little evidence of a response in aggregate consumer expenditures and concluded that consumers mostly saved the rebate. However, their aggregate analyses were soon overshadowed by the impressive household-level analysis.

Sahm, Shapiro, and Slemrod (2010, 2012) also estimated micro MPCs out of the 2008 rebate from rich survey data, but found MPCs around one-third, lower than those of Parker, Souleles, Johnson, and McClelland (2013). Noting that a significant part of Parker, Souleles, Johnson, and McClelland's (2013) high MPCs came from spending on motor vehicles, Sahm, Shapiro, and Slemrod (2012) calculated the implied fraction of aggregate motor vehicle sales that were induced by the 2008 rebate. They commented that this estimate was "surprisingly high" given that there were no dramatic shifts in motor vehicle sales around that time.<sup>1</sup> They cautioned, however, that their exercise did not allow for any partial or general equilibrium effects.

Most of the literature has overlooked Sahm, Shapiro, and Slemrod's (2012) important calculation. Figure I, which updates Ramey (2018), shows actual expenditures on new motor vehicles as the solid line, along with the implied counterfactual spending estimate depicted by the dashed line. This counterfactual is created as the difference between the actual spending and the estimated induced spending from the rebate using Parker, Souleles, Johnson, and McClelland's (2013) estimates.

The counterfactual implies that had there been no tax rebates, expenditures on motor vehicles would have declined from \$17 billion in March 2008 to less than \$3 billion in June 2008 and then would have rebounded sharply in late summer, averaging \$14

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1. See p. 242 and Table 14 of Sahm, Shapiro, and Slemrod (2012). Sahm, Shapiro, and Slemrod (2010) compare their own micro MPC estimates to total aggregate consumption in a similar exercise.

billion per month in August and September 2008. This counterfactual strains credulity, especially since the lowest actual level of motor vehicle expenditures during the Great Recession was \$12 billion in April 2009.

[ Figure I Here ]

In the first part of our paper, we extend the logic of the Sahm, Shapiro, and Slemrod (2012) exercise to a dynamic general equilibrium setting. We first construct a two-good, two-agent New Keynesian (TANK) model in which some households are life-cycle permanent income households and others are “hand-to-mouth” households who consume all their income. We calibrate the fraction of hand-to-mouth households in the economy and their dynamic spending propensities to match the Parker, Souleles, Johnson, and McClelland’s (2013) estimates. Aggregate consumption rises from both the direct micro effect of the rebate on household consumption and the induced macroeconomic effect through Keynesian multipliers. We then simulate the macroeconomic effects of a path of rebates that matches the timing and size of the actual 2008 rebate, which was announced in February and distributed mostly from April through July 2008. To create the counterfactual path of aggregate consumption in 2008 with no tax rebate, we subtract the model-simulated deviation from steady state from actual aggregate NIPA consumption.

The counterfactual paths imply that, in the absence of rebates, aggregate consumption would have declined steeply from May through July 2008 and recovered in August and September 2008, just as Lehman Brothers failed. The implied decline is larger than any other historical three-month decline, with the exception of the Covid-19 shutdown. Using narrative evidence and forecasts, we argue that this scenario is implausible.

Our claim about counterfactual aggregate consumption paths begs the question: How does one reconcile the high estimated micro MPCs from the literature with the implausible general equilibrium counterfactual paths? One possibility is an upward bias in the existing household MPC estimates. A second possibility is that general equilibrium forces, rather than magnifying the micro MPCs, actually dampen them. In the second half of the paper, we explore each of these possibilities and conclude that both are needed to explain the implausible counterfactuals.

To examine possible biases, we revisit Parker, Souleles, Johnson, and McClelland’s (2013) household estimates for the 2008 rebate using the Consumer Expenditure Sur-

vey (CEX). Their econometric methods were standard at the time, but subsequent econometric advances have discovered some biases in these methods. Kaplan and Violante (2014) raised important questions regarding the interpretation of the rebate coefficient as an MPC rather than a strict test of the permanent income hypothesis, and the subsequent econometric literature uncovered potential problems with event studies in general (e.g. Sun and Abraham (2020), Borusyak and Jaravel (2017), Borusyak, Jaravel, and Spiess (2024)). Building on their insights, we find that estimates of MPCs are affected by three separate biases: omitted variable bias, forbidden comparisons with previously treated households, and a rebate reporting bias. When we estimate a more general model that corrects these biases, the MPC estimates fall to 0.3. Moreover, when we distinguish between motor vehicles and all other spending, the MPC is 0.3 on motor vehicles and 0 on other consumption. Thus, we find no evidence in these data supporting the widely-used calibrations of an MPC of 0.25 or higher on nondurable goods.

Even with our lower micro MPC estimates, the baseline model still generates implausible macro counterfactuals because its general equilibrium forces amplify the micro MPCs. However, this model misses a key empirical fact: The relative price of motor vehicles spiked up when the 2008 rebates were distributed. The baseline model cannot capture this price movement because it follows the existing literature in assuming that durable and nondurable goods are perfect substitutes in production, implying a fixed relative price. To relax this restriction, we follow McKay and Wieland (2021) by allowing an upward-sloping relative supply curve, calibrated to micro evidence. Allowing for realistic movements in the relative price of motor vehicles generates substantial dampening in general equilibrium. The counterfactuals are no longer implausible because the aggregate consumption effects of the rebate are small: less than 20 cents for each dollar of rebates.

Three model elements are key to this general equilibrium dampening: (i) the majority of spending from the rebate is on motor vehicles; (ii) the short-run supply curve of motor vehicles is upward sloping; and (iii) motor vehicle demand is relatively elastic. Thus, the rebate-induced demand for motor vehicles from the hand-to-mouth households raises the relative price of motor vehicles and crowds out motor vehicle expenditure by optimizing households. The high demand elasticity implies that even a modest increase in the relative price can lead to substantial crowding out of durable expenditure by the optimizing households. This is true across a range of estimates for the vehicle de-

mand elasticity spanned by Bachmann, Born, Goldfayn-Frank, Kocharkov et al. (2021) and Baker, Kueng, McGranahan, and Melzer (2019).

Our findings imply that policy prescriptions from Heterogeneous Agent New Keynesian (HANK) models depend on the distribution of spending across nondurable and durable goods if the relative supply curve of durable goods is upward sloping in the short run. A total-spending micro MPC of 0.3, where all spending is concentrated on durables, leads to an aggregate consumption change of less than 0.1. In contrast, the same total spending MPC, where all spending is concentrated on nondurables, leads to an aggregate change of 0.4. Thus, a nondurable-only model with the same overall MPC predicts too large a stimulus from a tax rebate. This is because the intertemporal elasticity of substitution of nondurable demand is much less than for durable demand and the flat Phillips curve makes nondurable supply very elastic, resulting in no dampening.

The combination of dampening general equilibrium forces and more modest micro MPC estimates implies that the effect of the rebate on consumption expenditures in general equilibrium was modest. With our preferred micro MPCs of 0.3 on motor vehicles and 0 on nondurables, we find that the general equilibrium increase in total consumer spending was only 6 cents per dollar of the total rebate.

The paper is structured as follows. Section II provides a brief description of the details of the 2008 tax rebate and the behavior of aggregate disposable income and consumption in 2008. Section III presents the counterfactuals constructed from a standard two-agent, two-good New Keynesian model. It then argues that these counterfactual paths are implausible based on narratives, real-time forecasts, and comparisons with other historically-large drops in consumption.

The two subsequent sections reconcile the micro MPCs with the macro counterfactuals. Section IV revisits the micro estimates, demonstrates that they are biased upward, and presents new MPC estimates that are significantly lower. Section V modifies the model to incorporate more dampening effects in general equilibrium, which in conjunction with the smaller micro MPCs produces plausible macro counterfactuals. Section VI takes stock of the various findings and discusses the broader implications for using micro MPCs to calibrate macro models.

## II The 2008 Rebates

In February 2008, Congress and the Bush Administration enacted \$100 billion in tax rebates equal to eleven percent of January monthly disposable income. The amount of the rebate received by a household depended on tax status and dependents and was phased out at higher income levels; among the 85 percent of households receiving a check, the average amount was \$1,000. The timing of distribution was randomized according to the last two digits of the Social Security number.

As Figure II shows, almost half of the total was distributed in May, with most of the remainder distributed in June and July.

[ Figure II Here ]

Figure III shows the behavior of real NIPA disposable personal income and consumption expenditure from mid-2007 through mid-2009.<sup>2</sup> The vertical dashed line indicates May 2008 when almost half of the rebate checks were distributed. The effect of the 2008 tax rebate on disposable income is clearly evident in the spike in real disposable income series, but real consumption exhibits only a small bump.<sup>3</sup> These patterns in the aggregate data led Feldstein (2008) and Taylor (2009) to conclude that the aggregate impact of the rebate must have been small.

[ Figure III Here ]

One explanation for the modest consumption changes is that the BEA's smoothing procedures remove any high-frequency movements in personal consumption expenditures. We refute this argument in Appendix III.C: The underlying retail sales data features no spike in expenditure in summer 2008; well-measured components of expenditure such as car expenditure feature no spike; and the CEX data that Parker, Souleles, Johnson, and McClelland (2013) use to estimate high micro MPCs also features no

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2. For better illustration, real income and consumption are normalized to equal nominal values in January 2008 and the scaling of the y-axis is the same across the two graphs so that the variation in quantities can be compared.

3. Appendix III.B shows the behavior of the nominal series and discusses the behavior of inflation in 2008.

spike. In short, we find no role for measurement error in explaining the divergence of income and consumption expenditure in Figure III.

### III Macro Counterfactuals from a New Keynesian Model

In this section we use a standard medium-scale New Keynesian model, augmented with durable consumer goods, to show that the leading micro MPC estimates for the 2008 rebate lead to implausible counterfactuals for total consumption and motor vehicle expenditures even when we account for general equilibrium forces.<sup>4</sup> We first derive the model-implied counterfactuals and then use forecasts and comparisons with historical drops in consumption to argue that they are implausible.

#### III.A Two-Good, Two-Agent New Keynesian Model

Our macro model is a two-agent, two-good, medium-scale New Keynesian model. In the model, aggregate consumption rises due to both the direct micro effect of the rebate on consumption and the induced macro effect on income through Keynesian multipliers. We call the sum of these two effects on aggregate consumption the *general equilibrium marginal propensity to consume* of the rebate, or GE-MPC for short. We then subtract the simulated model deviations from steady state from actual aggregate consumption to create counterfactual paths of aggregate consumption in 2008, i.e., the path of aggregate consumption had there been no rebate.

We choose a two-agent model rather than a more general heterogeneous model for several reasons. First, the two-agent, two-good structure allows us to exactly target the intertemporal MPCs that are estimated from the micro data in 2008. In this way, our model will produce the identical increase in demand that we observe at the micro level even as we abstract from the complexity of micro foundations for hand-to-mouth behavior or durable expenditures. Second, the two-agent model is consistent with the negligible intertemporal MPCs after the first few months. Our new micro estimates using the CEX as well as other studies's estimates using higher frequency data find that all the spending from the 2008 rebates occurred within the first few weeks or couple months. For example, Table 4 of Sahm, Shapiro, and Slemrod (2010) shows that among

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4. Our approach is analogous to Chetty, Guren, Manoli, and Weber's (2013) demonstration that macro models calibrated with the large Frisch elasticities preferred by macroeconomists significantly over-predict the labor supply response to the Iceland tax holiday.

those households in the Michigan Survey who said they mostly spent the 2008 rebate, 36% spent it within a few weeks and an additional 50% spent it within three months. Borusyak, Jaravel, and Spiess (2024) revisit Broda and Parker’s (2014) estimates of the MPCs from the 2008 rebates using Nielsen data and find no impact on spending beyond two to four weeks (Figure 4C). In contrast, a distinguishing feature of Auclert, Rognlie, and Straub’s (2024) heterogeneous agent model is intertemporal MPCs that persist for several *years*.

Since the purpose of the model is to show how this increase in demand at the micro level gets propagated in general equilibrium, we consider a rich set of general equilibrium forces and calibrate their strength in accordance with the evidence. Thus, our approach builds on Auclert, Rognlie, and Straub (2024) and Wolf (2023), who show that the micro-level increase in demand and the strength of general equilibrium forces are sufficient statistics for how a demand shock gets propagated in general equilibrium, irrespective of the model structure that generates the increase in demand and the general equilibrium forces.

Our model is based on Ramey’s (2021) extension of Galí, López-Salido, and Vallés’s (2007) fiscal two-agent New Keynesian (TANK) model, but calibrated to a monthly frequency. The main addition to the model is a durable consumption good, which we interpret as motor vehicles. This part of the model builds on the recent analysis of durable goods expenditures by McKay and Wieland (2021, 2022).

We begin by describing the household’s problem in more detail since it is less standard than the other parts of the model. We then briefly summarize the other features, and refer interested readers to the appendix for more details.

*Optimizing Households.* A measure  $1 - \gamma$  of ex-ante identical households maximize utility subject to their budget constraints. Optimizing households form a family that provides consumption insurance across household members. To reduce the extremely high willingness to intertemporally substitute durables purchases that arises in standard models,<sup>5</sup> we assume that only a fraction  $1 - \theta^d$  of all optimizing households decide to reoptimize their durable stock at any point in time. This friction, which is motivated by Evans and Ramey’s (1992) model of calculation costs, produces a reversal in durable spending consistent with the evidence (e.g. McKay and Wieland (2021)) and keeps the model tractable because it produces a Calvo-type reduced form.<sup>6</sup>

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5. See Koby and Wolf (2020) and McKay and Wieland (2021).

6. In contrast, a conventional convex adjustment cost mechanically induces positive serial correlation in a household’s purchasing decisions and thereby overstates the crowding out from higher durable goods

The utility function for the family of optimizing households is:

$$(1) \quad E_0 \sum_{t=0}^{\infty} \beta^t \left[ \frac{(C_t^o)^{1-\frac{1}{\sigma}}}{1-\frac{1}{\sigma}} + \psi \frac{\int_0^1 D_t^o(i)^{1-\frac{1}{\sigma^d}} di}{1-\frac{1}{\sigma^d}} - \nu \frac{(H_t^o)^{1+\phi}}{1+\phi} \right]$$

where  $C_t^o$  is nondurable consumption,  $D_t^o(i)$  is the durable stock of household  $i$ , and  $H_t^o$  is hours worked.<sup>7</sup> For brevity, only the durables stock is indexed by household  $i$  since the other arguments are identical across households. The aggregate household budget constraint is

$$(2) \quad A_t^o = \frac{R_{t-1}}{\Pi_t} A_{t-1}^o - C_t^o + W_t H_t^o - X_t^o - OC_t^o - T_t^o + \text{Profits}_t^k + \text{Profits}_t^s$$

$$(3) \quad X_t^o = p_t^d \left[ \int_0^1 [D_t^o(i) - (1 - \delta^d) D_{t-1}^o(i)] di \right]$$

$$(4) \quad OC_t^o = \eta \int_0^1 D_t^o(i) di$$

where  $R_t$  is the gross nominal interest rate,  $\Pi_t$  is the gross inflation rate measured in nondurable goods prices,  $A_t^o$  are holdings of the nominal bond,  $W_t$  is the real wage,  $T_t^o$  are net taxes (i.e. taxes less transfers),  $\text{Profits}_t^k$  are profits of the capital good producing firms, and  $\text{Profits}_t^s$  are profits of the sticky-price firms, which produce nondurable goods.  $X_t^o$  is net durable expenditures denominated in nondurable goods, and are the sum of net durable purchases of each household,  $D_t^o(i) - (1 - \delta^d) D_{t-1}^o(i)$ .  $OC_t$  are operating costs for the durable good (e.g., gasoline) which are a fraction  $\eta$  of the total durable stock held by all households. The inclusion of operating expenditures also helps produce more realistic elasticities of durable demand.

Optimizing households pick an optimal plan  $\{C_t^o, A_t^o, D_t^o(i)\}_{t=0}^{\infty}$  to maximize utility. Labor supply is not chosen by the household, but instead by a union as discussed below.

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prices. For richer models of household durable decisions, see, for example, Carroll and Dunn's (1997) and Attanasio, Larkin, Ravn, and Padula's (2022) household-level analysis or McKay and Wieland's (2021) general equilibrium analysis.

7. We assume separable utility for tractability of the durable choice problem subject to a Calvo friction. We have explored specifications in which we can directly specify the elasticity of substitution for durable goods. These results are quantitatively similar to our baseline case and are available upon request.

The first order conditions for  $C_t^o, A_t^o$  are:

$$(5) \quad \lambda_t = (C_t^o)^{-\frac{1}{\sigma}}$$

$$(6) \quad \lambda_t = \beta \frac{R_t}{\Pi_{t+1}} \lambda_{t+1}$$

where  $\lambda$  is the Lagrange multiplier on the household budget constraint.

We next derive the optimal choice of  $D_t^o(i)$  conditional on making an adjustment. Because the durable stock of household  $i$  in the problem is separable from the durable stock of other households, the durable part of the optimization problem for household  $i$  is simply,

$$(7) \quad \max_{D_t(i)} \sum_{s=0}^{\infty} (\beta \theta^d)^s \left[ \psi \frac{[(1-\delta^d)^s D_t(i)]^{\frac{\sigma^d-1}{\sigma^d}}}{1-\frac{1}{\sigma^d}} - \lambda_{t+s} \eta (1-\delta^d)^s D_t(i) \right] - \lambda_t p_t^d D_t(i) \\ + \sum_{s=0}^{\infty} \beta^s (\theta^d)^{s-1} (1-\theta^d) \lambda_{t+s} p_{t+s}^d (1-\delta^d)^s D_t(i)$$

Here  $(\theta^d)^s$  is the survival probability of the current durable stock into period  $s$ ,  $\psi \frac{D_t(i)^{\frac{\sigma^d-1}{\sigma^d}}}{1-\frac{1}{\sigma^d}}$  is its contribution to household utility,  $\lambda_{t+s} \eta D_t(i)$  is the operating cost while the durable stock remains in place, measured in utils,  $\lambda_t p_t^d D_t(i)$  is the purchasing price in utils, and  $\lambda_{t+s} p_{t+s}^d (1-\delta^d)^s D_t(i)$  is the resale value of the durable in utils if another adjustment opportunity arises at time  $t+s$ .

The problem is identical across households that can make an adjustment at time  $t$ . Therefore, let  $D_t^{o*}$  denote the optimal reset value for the durable stock at time  $t$ . In Appendix A we show that the first order conditions of the problem can be written as,

$$(8) \quad D_t^{o*} = \left( \frac{\Omega_{1t}}{\Omega_{2t}} \right)^{\sigma^d}$$

$$(9) \quad \Omega_{1t} = \psi + \beta \theta^d (1-\delta^d) \Omega_{1,t+1}$$

$$(10) \quad \Omega_{2t} = (p_t^d + \eta) \lambda_t - \beta (1-\delta^d) p_{t+1}^d \lambda_{t+1} + \beta \theta^d (1-\delta^d) \Omega_{2,t+1}$$

$\Omega_1$  is the expected present discounted value of a unit of durable varieties and  $\Omega_2$  is the expected present discounted value of the user cost. The last two equations express the  $\Omega$ 's recursively.

By defining the total durable stock as

$$(11) \quad D_t^o \equiv \int_0^1 D_t^o(i) di,$$

we obtain the standard durable accumulation equation and durable net expenditure as a function of aggregate variables only,

$$(12) \quad D_t^o = (1 - \delta^d)D_{t-1}^o + \frac{X_t^o}{p_t^d}$$

$$(13) \quad X_t^o = p_t^d(1 - \theta^d)[D_t^{o*} - (1 - \delta^d)D_{t-1}^o].$$

$D_t^{o*}$  is the optimal stock of durables for households that adjust. The expression for durable purchases shows that the calculation cost friction directly limits the extensive margin of durable adjustment to  $(1 - \theta^d)$ . In Appendix A we show that the friction also limits the sensitivity of the intensive margin—the term in brackets—to the real interest rate.

*Hand-to-Mouth Households.* In order for lump-sum transfers to have general equilibrium effects, we require non-Ricardian households. We adopt Galí, López-Salido, and Vallés's (2007) assumption that a certain fraction  $\gamma$  consume hand-to-mouth. Relative to their set-up, our hand-to-mouth households may consume their income over several periods rather than all at once.

We assume that in steady state, hand-to-mouth households have the same after-tax income and consume the same relative amount of durable and nondurable services as optimizing households,

$$(14) \quad WH^m - T^m = WH^o - T^o + \text{Profits}^k + \text{Profits}^s$$

$$(15) \quad \frac{C^m}{X^m} = \frac{C^o}{X^o}$$

where variables superscripted by  $m$  denote the hand-to-mouth household.

We then directly specify dynamic marginal propensities to consume for nondurable and durable expenditures to match both the allocation across goods and any lagged

effects implied by the micro MPC estimates,

$$\begin{aligned}
(16) \quad C_t^m - C^m + \eta(D_t^m - D^m) &= \sum_{l=0}^L mpc_l [W_{t-l} H_{t-l}^m - T_{t-l}^m - (WH^m - T^m)] \prod_{k=1}^l \frac{R_{t-k}}{\Pi_{t-k+1}} \\
X_t^m - X^m &= \sum_{l=0}^L mpx_l [W_{t-l} H_{t-l}^m - T_{t-l}^m - (WH^m - T^m)] \prod_{k=1}^l \frac{R_{t-k}}{\Pi_{t-k+1}} \\
1 &= \sum_{l=0}^L (mpc_l + mpx_l) \\
mpx_l &= \frac{\vartheta}{1 - \vartheta} mpc_l, \quad \forall l = 0, \dots, L
\end{aligned}$$

where  $mpc_l$  is the marginal propensity to spend on nondurable goods today out of income  $l$  periods ago, and  $mpx_l$  is the marginal propensity to spend on durable goods today out of income  $l$  periods ago. Income that was saved  $l$  periods ago for consumption today accrues real interest  $\prod_{k=1}^l \frac{R_{t-k}}{\Pi_{t-k+1}}$ .

*Overview of Other Features* The remaining elements of the model are standard. Intermediate goods firms are monopolistically competitive and face a Calvo-style adjustment cost on prices. Intermediate goods can be turned one-to-one into either nondurable goods or durable goods, which implies their relative price is constant and equal to  $p_t^d = 1$ . In labor markets, unions mark up nominal wages over the marginal rate of substitution and face Calvo-type adjustment costs. The result is that short-run employment fluctuations are driven more by labor demand than labor supply. Firms face an adjustment cost on capital investment, but they can vary their utilization of capital, so capital services are more cyclical than the capital stock. The result is more elastic output supply since it mutes the diminishing returns to labor and prevents real marginal cost from increasing much when output rises. The monetary rule is inertial, with a long-run coefficient of 1.5 on the inflation gap and 1/12 on the monthly output gap. Lump-sum taxes respond to the deviation of government debt from its steady-state values but with a lag of one year. A more complete description with equations is provided in the appendix.

[ Table I Here ]

### III.B Calibration

The calibrated parameters with their descriptions are shown in Table I. Note that the model is calibrated to a monthly frequency. In addition to the calibrations shown in the table, we calibrate the steady-state transfers by type of household so that hand-to-mouth and life-cycle permanent income households consume the same amount in the steady state. The durable goods parameters are chosen to match the average share of motor vehicle spending in PCE and its depreciation rate in the fixed asset table. Operating costs are based on PCE expenditures on motor vehicle fuels, lubricants, and fluids. The appendix shows more details of the model.

The timing of spending by hand-to-mouth households is important for constructing the counterfactual path of consumption. We assume that the hand-to-mouth households respond to a shock to their disposable income by spreading their spending over three months. Estimates from Broda and Parker (2014) using higher-frequency Nielsen data on nondurable expenditures suggest that two-thirds of expenditure occurs in the month of the rebate, and one-sixth each in the following two months.<sup>8</sup> In our own investigation using CEX data, we find no evidence of additional expenditure after three months.<sup>9</sup> Unfortunately, the CEX does not lend itself to estimate monthly expenditure patterns as most households report expenditures divided equally across the three months within an interview. One exception to this limitation is reported car expenditure, which more precisely identifies the month of purchase. Appendix Table C.8 shows that the car expenditure response occurs in the three months around the rebate. We conservatively choose an equal spread of expenditure since this minimizes the extent of V-shapes in our counterfactuals and is thus more consistent with larger MPCs.

We simulate several versions of the model, across a range of fractions of households who are hand to mouth. We set values for  $\gamma$ , and thus a three-month cumulative MPC, equal to 0.3, 0.5, and 0.9. The lower value, 0.3, reflects our preferred estimate based on our new estimates that correct for several biases (presented in the next section in Table III, column 4). The other two values, 0.5 and 0.9, are the estimates from our replication of Parker, Souleles, Johnson, and McClelland (2013) in the full CEX sample and the subsample of rebate recipients (Appendix Table C.6, column 1).

A key distinction in both the estimates and in our model is the allocation of spending between nondurable goods and motor vehicles. We again calibrate these to empirical

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8. Borusyak, Jaravel, and Spiess (2024) also do not find evidence of spending after three months.

9. See the implied 6-month MPC in Table III, column 4.

estimates. In our preferred specification the MPC on motor vehicles is 0.3 (Table V). Using the Parker, Souleles, Johnson, and McClelland (2013) specifications we obtain an MPC on motor vehicles of 0.4 in the full CEX sample and the subsample of rebate recipients (Table C.7, column 1).

The curvature of durable utility  $\sigma^d$  and the Calvo durable good adjustment probability  $\theta^d$  determine how sensitive durable demand is to general equilibrium changes in durable prices and the real interest rate. We calibrate these parameters based on estimates of the demand elasticity at the household level, which difference out any local or aggregate general equilibrium price effects.

First, we set the long-run demand elasticity for vehicles to  $\sigma^d = -1$  based on an average of three existing household studies.<sup>10</sup> Second, we calibrate the durable Calvo probability  $\theta^d$  to target an increase in durable demand of 15% over six months in anticipation of a 1% increase in prices, as estimated by Bachmann, Born, Goldfayn-Frank, Kocharkov et al. (2021).<sup>11</sup> The implied parameter value for  $\theta^d$  varies across values of the fraction of hand-to-mouth consumers since these do not respond to intertemporal price changes. For example when we target an MPC of 0.3, then we obtain  $\theta^d = 0.85$ .

### III.C General Equilibrium Counterfactuals

With the model constructed and calibrated, we now compute counterfactual paths of consumption that take into account the full dynamic general equilibrium effects. We start the economy in steady state in January 2008, and assume that households do not anticipate the equilibrium path of prices resulting from the rebate until after the first rebate payments are made in April.<sup>12</sup> We feed a path of rebate shocks into the model that matches the relative size and timing of the actual rebate shown in Figure II.

We use first-order perturbation methods to solve for the general equilibrium impulse responses of the variables to the path of rebates. We then construct macro-

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10. McCarthy (1996), Bento, Goulder, Jacobsen, and Von Haefen (2009), and Dou and Linn (2020).

11. Table A.4., columns 3 and 9. Our calibrated demand elasticity is consistent with Mian and Sufi (2012) estimates for the 2009 Cash-for-Clunkers program. See Appendix III.E for details.

12. Without this assumption, optimizing households would foresee the future rise in the real interest rate and increase their purchases of durable goods immediately. Because the real interest rate moves very little in the baseline model, this assumption is not very important in this section. However, this assumption is important in Section V because there the relative price of motor vehicles moves in equilibrium. In that section we also show that the relative motor vehicle price does not rise until May 2008 (Figure VII), and that motor vehicle expenditure by non-recipient households drops in May 2008 and that there is no unusual increase in motor vehicle expenditures earlier in the sample (Figure C.8). These outcomes are consistent with our assumption that households do not anticipate the effects of the rebate on prices.

counterfactuals by subtracting the model-implied impulse response functions for consumer expenditures from the observed consumer expenditure data. Because the model is linearized, the counterfactuals for the tax rebate would be identical if we also fed the model with other shocks that hit the economy at the time.

[ Figure IV Here ]

Figure IV plots counterfactual total consumption and motor vehicle expenditure paths based on both the micro MPCs, which exclude any general equilibrium effects, and the GE-MPCs, which incorporate full dynamic general equilibrium feedbacks. The counterfactuals in the top left panel that do not allow for general equilibrium effects are the analogs to the Sahm, Shapiro, and Slemrod (2012) exercise we showed in the introduction. The figures show prominent, and we will argue implausible, V-shapes for total consumption. According to these counterfactuals, consumption would have collapsed from May through July 2008 and recovered in August 2008 before beginning the longer downward path starting with the fall of Lehman Brothers.

The top right panel of Figure IV shows that allowing for general equilibrium effects makes the counterfactual even more V-shaped. The highly transitory nature of the rebate coupled with a flat Phillips curve and interest rate inertia implies that there is little crowding out through the real interest rate.<sup>13</sup> The dominant general equilibrium force is the Keynesian multiplier. Thus, the effects of the rebate are amplified in general equilibrium, particularly as the micro MPCs become larger, so the counterfactual paths become even more V-shaped.

The bottom two panels show the counterfactuals for real motor vehicle expenditure. The left panel only accounts for the direct effect of the rebate, excluding any general equilibrium effects. This exercise is similar to Sahm, Shapiro, and Slemrod (2012) except that it accounts for all motor vehicle expenditure, not just new cars.

The right panel includes all general equilibrium effects. The V-shapes of counterfactual motor vehicle expenditure are more even more pronounced than for total PCE.

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13. In our calibrated Taylor Rule, the central bank raises the nominal rate by only 17 to 56 basis points for micro MPCs ranging from 0.5 to 0.9. Complete offset of the effects of the rebate would require nominal interest rate hikes of 6.6 to 10 percentage points in June 2008. This increase is not only implausibly large but also implies infeasible negative nominal rates in the counterfactual since the actual Federal Funds Rate stood at 2.5%. Thus, our model implies that monetary policy cannot account for the divergence between large micro MPCs and the apparent lack of movement in expenditure at the aggregate level.

The counterfactual drop in motor vehicle expenditure from April to July ranges from 40% to 70% across the range of MPCs. This reflects that the MPCs on motor vehicles in the micro estimates are large relative to the overall size of motor vehicle expenditures in consumption expenditure, and that these direct expenditures get further amplified in general equilibrium.

### **III.D Assessing the Plausibility of the Baseline Counterfactuals**

We now use narrative evidence, forecasts, and comparisons with historical consumption drops to argue that all three counterfactuals shown in the last section are implausible. We show that none of the events at that time would have led aggregate consumption to fall dramatically in spring and summer 2008 and then recover just as Lehman Brothers was failing.

January 2008 began with negative economic news. The employment report for the previous December showed a jump in the unemployment rate, leading forecasters and policymakers to worry that a recession was imminent. In response, the Federal Reserve began lowering interest rates in January and Congress and Administration enacted the rebates in February.

Goldman Sachs released their forecast in early January 2008 and were among the first to predict that the U.S. was already in recession. Their forecasts were based on the following assumptions.<sup>14</sup> First, the Fed would cut the federal funds rate target from 4.25 to 2.5 by the end of the year, with the first 50 basis point cut at the next FOMC meeting on January 30th. Second, housing prices would decrease 20 to 25 percent below their peak. Third, Congress and the President would pass a temporary tax break as part of a fiscal stimulus plan later in the year.

Goldman Sachs forecasted no change in real consumption expenditures (PCE) in 2008Q1, a decrease of 0.125 percent (not annualized) in each of 2008Q2 and 2008Q3, and a 0.25 percent increase in 2008Q4. Thus, they forecasted actual declines in real consumption expenditures, but they were tiny in magnitude. Similarly, contemporary forecasts from the Federal Reserve Board Staff (Greenbooks) and the Survey of Professional Forecasters did not predict large drops of consumption in summer 2008. Most forecasters predicted an increase in real consumption and even the most pessimistic

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14. This summary is based on contemporaneous news accounts, such as the CNN Money article "Recession may already be here," January 10, 2008.

forecaster from the Survey of Professional Forecasters ("SPF Min.") predicted only a small decrease in consumption in summer 2008. All these forecasts are shown alongside actual values in the left panel of Figure V.<sup>15</sup> None predicted that consumption would decline significantly in Summer 2008.

[ Figure V Here ]

However, the forecasters in January 2008 were not certain the economy was in recession and they did not foresee the rapid run-up in oil prices in spring and summer or the Lehman Brothers failure in September. Crude oil prices rose from \$98 per barrel in January 2008 to a peak of \$140 per barrel in July 2008 and then fell to \$33 per barrel by the end of the year. All these factors could have negatively affected consumption. Thus, we construct our own forecasts that factor in those negative events to create more pessimistic forecasts to compare to our counterfactuals.

Our forecasting model is a monthly-frequency time series model with current and six lags of the following endogenous variables: log real consumption, log real disposable income, log consumption deflator, and the Gilchrist and Zakrajšek (2012) excess bond premium.<sup>16</sup> We estimate two versions of the model. The pessimistic model conditions on the actual path of oil prices, a dummy variable for recession, and a dummy variable for the Lehman Brothers bankruptcy in September 2008. Current values and six lags are included in the model. The regular model excludes the recession and Lehman dummies and assumes that real oil prices are endogenous, which results in real oil prices remaining roughly constant in the forecast. We estimate the models on data from 1984m1 - 2019m12 and forecast dynamically starting in January 2008 before the rebates were passed.<sup>17</sup> In Appendix Figure C.7 we show that forecasts based on other variable combinations generally lie between the regular and the pessimistic forecast. The figure also shows the confidence intervals for our pessimistic forecast.

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15. In each case, we select the last survey prior to the passage of the Economic Stimulus Act of 2008 since afterward forecasters would include the rebate response as part of their forecast. The January Greenbook actually does incorporate the tax rebates in their consumption forecasts and assumes an MPC of 0.67 (Edelberg and Sahm, 2008). However, they forecasted that the rebates would be received in the second half of 2008, not in the second quarter when most rebates were actually sent.

16. We explored the addition of a number of other variables, such as consumer confidence, but they did not noticeably change the forecasts and/or were not statistically significant.

17. We start the estimation period in 1984 because we found evidence of structural breaks, likely due to the Great Moderation and the change in the effects of oil prices on consumption expenditures (e.g. Edelstein and Kilian (2009)).

The right panel of Figure V shows the actual data and the monthly-frequency forecasts. The regular forecast is similar to the most pessimistic projections of the professional forecasters shown in the left panel. In contrast, our pessimistic forecast predicts noticeably less consumption. The forecast lies below the actual total consumption path from April 2008 through October 2008, which is consistent with some stimulus effect of the rebates. However, the cumulative difference between actual consumption and the most pessimistic forecast is only \$20 billion. With a total rebate of \$100 billion, the implied GE-MPC is only 0.2 even when we attribute the entire difference to the effects of the rebate.

Figure IV from the last section shows that this pessimistic forecast (denoted with a dashed line) lies above all three counterfactuals in the summer of 2008 and does not exhibit any V-shape. Thus, even our pessimistic forecast does not predict the sharp decline in consumption implied by the counterfactuals.

[Table II Here]

Table II puts those counterfactual declines into further perspective by comparing them to other episodes since 1959. The counterfactuals predict a 5.9% decline in PCE from April through July 2008 for a micro MPC of 0.9 and a 2.6% decline in PCE for a micro MPC of 0.5. Such sharp declines over a three-month window are exceedingly rare. Only the COVID-19 lockdowns caused a larger drop in PCE. And only another extreme macroeconomic event — the 1980 Volcker disinflation coupled with credit controls — generates a PCE drop of comparable magnitude to the counterfactual with an MPC of 0.5. The next two largest historical declines in 1960 and 1974 followed anomalous large upward spikes in consumption expenditures.<sup>18</sup> The last line of Panel A shows that the failure of Lehman brothers did not produce such sharp declines in consumption expenditures.

Panel B of Table II shows that the implied counterfactuals for motor vehicle expenditures are outside any historical experience since monthly expenditure data became available. With the Parker, Souleles, Johnson, and McClelland (2013) micro MPCs the model counterfactual predicts a three-month drop in motor vehicle expenditures from

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18. The August 1974 spike occurred when U.S. auto manufacturers announced dramatic price increases for the 1975 model year. In response, consumers rushed to buy the 1974 model year autos before they sold out. We could not determine the source of the anomalous spike in 1960.

50% to 70%. Yet the largest observed decline since 1959 is 30% during the COVID shutdown. Even with a micro MPC of 0.3, the model predicts that, absent the rebate, the summer of 2008 would have been worse than the COVID shutdown and followed by a swift economic recovery.

In short, the counterfactuals imply that the macroeconomy was under extreme stress in the summer of 2008 and only the rebate prevented a decline in consumer expenditures of historic proportions. But whatever caused this stress was short-lived since the counterfactuals show that the economy would have swiftly recovered in a V-shaped manner. Yet, there were no signs of such extreme, short-lived macroeconomic stress in anyone's forecasts. In contrast, historical declines in expenditures similar or smaller than the model counterfactuals are associated with clear macroeconomic events such as COVID-19 and the onset of credit controls in conjunction with the Volcker disinflation. This suggests that the counterfactuals are not plausible.

We draw two main conclusions from this section. First, a comparison of the actual behavior of aggregate consumption with our most pessimistic forecast suggests that the GE-MPC out of the rebate must have been 0.2 or less. Second, this aggregate GE-MPC is inconsistent with the implications of a standard one-good NK model calibrated with micro MPCs above 0.2 for total expenditures. Thus, either the existing micro MPC estimates are biased upwards or the standard NK model is missing important GE dampening. In the next two sections, we show that both statements are true.

## **IV Revisiting the 2008 Rebate Micro MPC Estimates**

This section provides the first part of our reconciliation of the micro MPC estimates and the macro counterfactuals. We revisit the Parker, Souleles, Johnson, and McClelland (2013) micro MPC estimates from the Consumer Expenditure Survey and show that they are biased upward. After correcting for the biases, we estimate the micro MPC is around 0.3, with almost all the spending on durable goods. Since even these estimates lead to implausible macro counterfactuals in our previous model, we complete the reconciliation in the subsequent section by showing how modifying our macro model to make the supply of durable goods less elastic results in general equilibrium forces that dampen rather than amplify the micro MPCs.

The most widely cited micro MPC estimates, which range from 0.5 to 0.9, come from Parker, Souleles, Johnson, and McClelland (2013). The authors worked with the

U.S. Bureau of Labor Statistics to add a question about the date of receipt of the 2008 rebate to the monthly Consumer Expenditure Survey (CEX), a rotating panel survey of household expenditure. Furthermore, since rebate checks were sent to households based on the last two-digits of their social security number, the timing of treatment (i.e. distribution of the rebate) was effectively random. Parker, Souleles, Johnson, and McClelland (2013) leverage the variation in treatment time (i.e., the month in which the household received the rebate) and in some cases the treatment size (i.e. the dollar value of the rebate check) to estimate the causal impact of receiving a rebate on household spending using a standard difference-in-differences (DID) event-study methodology.

In this section, we document and correct for three important upward biases in the Parker, Souleles, Johnson, and McClelland (2013) estimation method: (1) An omitted variable bias from not allowing for lagged rebate effects; (2) a bias from “forbidden comparisons” across cohorts with heterogeneous treatment effects; and (3) a rebate reporting bias stemming from a correlation between lagged expenditure and the report of receipt of a rebate. When we correct for these biases, we estimate substantially reduced micro MPCs of around 0.3.

Our econometric analysis builds and expands on questions raised a decade ago by Kaplan and Violante (2014). They noted that the coefficient on the rebate in Parker, Souleles, Johnson, and McClelland’s (2013) specification cannot be interpreted as an MPC because it omits the lagged effect of the rebate on changes in consumption. Our analysis also builds on and complements the work of Borusyak and Jaravel (2017) and Borusyak, Jaravel, and Spiess (2024) highlighting the problem of “forbidden comparisons” in event studies. They apply their new method to the MPC estimates of Broda and Parker (2014) using Nielsen data and also find a sharp reduction in the MPC estimates. We show that we obtain similar reductions in MPC estimates when we use their method on the CEX data.

#### **IVA Baseline Parker, Souleles, Johnson, and McClelland (2013) Specification and Replication**

To estimate the causal impact of receiving a check on household consumption, Parker, Souleles, Johnson, and McClelland (2013) estimate several versions of a standard regression used for testing the permanent income hypothesis. The version we

focus on is,

$$(17) \quad C_{i,t} - C_{i,t-1} = \sum_s \beta_{0s} month_s + \beta_1' X_{i,t} + \beta_2 I(ESP_{i,t}) + u_{i,t}$$

where  $t$  indexes the interview, and  $i$  indexes individual households. In each interview, respondents are asked to report their expenditure  $C_{it}$  over the three months prior to the interview month. The interview is repeated every three months for four times. The regression includes fixed effects for each month ( $month_s$ ), household controls for age and change in household size  $X_{i,t}$ , and the main variable of interest,  $I(ESP)$ , which is a dummy variable equal to one if the household received a rebate, i.e., an Economic Stimulus Payment (ESP).

We make two changes to the original Parker, Souleles, Johnson, and McClelland (2013) specification for the purposes of our analysis: First, we estimate MPCs for total expenditure using the BEA definitions for PCE because we construct counterfactuals for PCE.<sup>19</sup> The biggest change relative to total expenditure in Parker, Souleles, Johnson, and McClelland (2013) is that our estimates net out sales of used vehicles. Second, we drop all households that report receiving a rebate in more than one interview as multiple instances of treatment complicate the interpretation of  $\beta_2$  as a micro MPC for a single income change.

Columns (1) of Table III reports the estimates for  $\beta_2$  from equation (17). Panel A reports the estimates the treatment effects for the full sample and Panel B for the rebate-only sample. The full sample has more power because households that receive rebates are also being compared to households that never receive the rebate. For this comparison to be valid these groups of households must be on parallel trends. The rebate-only sample does not require this assumption as it only makes comparisons among households that report receiving a rebate, but this comes at the cost of statistical precision.

The estimates in column (1) align closely with Parker, Souleles, Johnson, and McClelland (2013). We estimate a \$470 response in the full sample (Panel A), compared to \$495 in Parker, Souleles, Johnson, and McClelland (2013). Appendix Table C.5 reports the corresponding rebate income, \$950, which implies an MPC of 0.5. Parker, Souleles, Johnson, and McClelland (2013) do not report a dollar response for the rebate-only

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19. We map the CEX UCC codes into PCE categories using the concordance provided by the BLS: <https://www.bls.gov/cex/cepceconcordance.htm>

sample but our MPC of above 0.8 in column (1) of Panel B is again very close to their value.

We next show how that these MPC estimates are upward biased.

#### IV.B Bias from Omitting the Lagged Rebate

The first bias we identify is an omitted variable bias owing to serial correlation in the treatment variable. To understand this bias, suppose the true model for consumer expenditure is

$$(18) \quad C_{it} = \alpha_i + \lambda_t + \beta I(ESP_{i,t}) + \epsilon_{it}$$

where  $\alpha_i$  and  $\lambda_t$  are fixed effects and  $I(ESP_{i,t})$  is a treatment indicator equal to 1 when the household receives a rebate. We assume that the timing of the treatment is random and that households are treated only once.

[Table III Here]

To align with the baseline specification (17), we take first differences,

$$(19) \quad \Delta C_{it} = \Delta \lambda_t + \beta I(ESP_{i,t}) + \eta_{it}, \quad \text{with} \quad \eta_{it} \equiv -\beta I(ESP_{i,t-1}) + \Delta \epsilon_{it}.$$

Thus, Parker, Souleles, Johnson, and McClelland's (2013) first-difference specification includes the lagged rebate indicator  $I(ESP_{i,t-1})$  in the error term  $\eta_{it}$ .

To assess the resulting bias, first define  $\hat{X}_t$  as the residual from regressing a variable  $X_t$  on a time fixed effect. Then the OLS estimator for the contemporaneous rebate effect can be written as,

$$(20) \quad \begin{aligned} \beta^{OLS} &= \frac{\text{Cov}\left(\widehat{\Delta C}_{it}, \widehat{I(ESP_{i,t})}\right)}{\text{Var}\left(\widehat{I(ESP_{i,t})}\right)} \\ &= \beta - \beta \frac{\text{Cov}\left(\widehat{I(ESP_{i,t-1})}, \widehat{I(ESP_{i,t})}\right)}{\text{Var}\left(\widehat{I(ESP_{i,t})}\right)}. \end{aligned}$$

The covariance,  $Cov\left(\widehat{I(ESP_{i,t-1})}, \widehat{I(ESP_{i,t})}\right)$ , is negative in a setting with staggered treatment because current treatment reduces the probability of treatment in the following period. When the treatment effect is positive ( $\beta > 0$ ),  $\beta^{OLS}$  is upward-biased. Intuitively, households treated at  $t$  are being compared to households treated at  $t - 1$ , whose consumption is falling as the effect of the rebate on the level of consumption reverses. This contaminated control inflates the OLS estimate of  $\beta$ . Note that  $\beta^{OLS}$  is unbiased under the null of the Permanent Income Hypothesis (PIH)  $\beta = 0$ .<sup>20</sup> Thus, Equation (17) is a valid test of the PIH, but the point estimates for  $\beta_2$  cannot be interpreted as MPCs, as previously shown by Kaplan and Violante (2014).

To show the importance of this bias in our setting, Figure VI plots the period-by-period treatment effects that make up the total treatment effect  $\beta_2$  in (17). Following Sun (2021) we decompose each period treatment effect into two parts: The contribution from comparing rebate recipients with households that have not yet or will never receive a rebate (dark bars) and the contribution from comparing rebate recipients with households that have previously received a rebate (light bars). Due to the three-month rotating panel structure of the CEX and the first rebates being reported in June, the first comparisons with previously treated households are made in September. The light bars for September show that these comparisons imply very large positive treatment effects — \$670 in the full sample and \$2200 in the rebate only sample. But this effect may simply reflect mean-reversion of the June cohort rather than a treatment effect for the September cohort.

[ Figure VI Here ]

To determine how much the estimated propensity to spend in the Parker, Souleles, Johnson, and McClelland (2013) equation is inflated by mean-reversion of previously treated units, we estimate an alternative model in which we add a rebate lag to equation (17),

$$(21) \quad C_{i,t} - C_{i,t-1} = \sum_s \beta_{0,s} month_s + \beta_1' \mathbf{X}_{i,t} + \beta_2 I(ESP_{i,t}) + \beta_3 I(ESP_{i,t-1}) + u_{i,t}$$

Column (2) of Table III reports estimates of the contemporaneous effect  $\beta_2$  and the lagged effect  $\beta_3$ . The contemporaneous spending effect shrinks by \$40 in the full sam-

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20. The PIH predicts  $\beta = 0$  because the change in income from the rebate is anticipated.

ple, indicating that the original estimates were upward biased. In the rebate-only sample the contemporaneous effect of the rebate falls by almost \$240. In both samples the estimate on the lagged rebate coefficient is negative, consistent with spending reversals causing an upward bias when the lagged rebate variable is omitted. The fact that the bias is more severe in the rebate-only sample is expected since relatively more variation in this sample comes from comparing rebate recipients to previously treated households.<sup>21</sup>

#### IV.C Heterogeneous Treatment Effects and Forbidden Comparisons

The lagged rebate indicator in (21) will account for the typical mean-reversion of consumer expenditure after receiving a rebate. However, Figure VI shows that the treatment effects of the rebate may vary substantially by date of receipt. For example, in the full sample the propensity to spend is particularly large for the June cohort. We would therefore expect greater mean-reversion for the June cohort than the July cohort. But  $\beta_3$  in (21) will only account for the average mean-reversion, not for the likely larger mean-reversion of the June cohort. Thus, the comparison of the September cohort with the June cohort after accounting for average mean-reversion may still be contaminated by lagged treatment effects.<sup>22</sup>

Formally, suppose the true model for consumer expenditure is

$$(22) \quad C_{it} = \alpha_i + \lambda_t + \beta^i I(ESP_{i,t}) + \epsilon_{it}$$

where the rebate effect  $\beta^i$  may now differ across individuals. First differencing to align with the Parker, Souleles, Johnson, and McClelland (2013) specification, the equation becomes

$$(23) \quad \begin{aligned} \Delta C_{it} &= \Delta \lambda_t + \beta I(ESP_{i,t}) + \gamma I(ESP_{i,t-1}) + \eta_{it}, & \gamma &\equiv -\beta, \\ \eta_{it} &\equiv (\beta^i - \beta) I(ESP_{i,t}) - (\beta^i - \beta) I(ESP_{i,t-1}) + \Delta \epsilon_{it} \end{aligned}$$

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21. In their Table 5, Parker, Souleles, Johnson, and McClelland (2013) report estimates from a specification with a lagged rebate variable. Our estimates in column 2 of Table III are consistent with theirs as they also find that the magnitude of the estimate of  $\beta_2$  declines. But their discussion focuses on the long-run estimates of MPCs implied by this specification, rather than correcting for an omitted variable bias.

22. We do not need to take a stand on the source of treatment effect heterogeneity. Even if it just reflects sampling noise, the OLS estimates will not recover an average treatment effect.

and the OLS estimator for the contemporaneous rebate effect is,

$$(24) \quad \beta^{OLS} = \beta + \frac{\text{Cov}(\widehat{\beta_i I(ESP_{i,t})} - \beta I(ESP_{i,t}), \widehat{I(ESP_{i,t})})}{\text{Var}(I(ESP_{i,t}))} - \frac{\text{Cov}(\widehat{\beta_i I(ESP_{i,t-1})} - \beta I(ESP_{i,t-1}), \widehat{I(ESP_{i,t})})}{\text{Var}(I(ESP_{i,t}))}$$

where  $\widehat{I(ESP_{i,t})}$  is the residual from the regression of  $I(ESP_{i,t})$  on a time fixed effect and  $I(ESP_{i,t-1})$ . The last covariance represents possible contamination bias from using later treated groups to correct for the spending reversal of earlier treated groups. The first covariance captures that OLS will put relatively more weight on the earlier treated group because there is more variation in the purified treatment  $\widehat{I(ESP_{i,t})}$ .

A simpler expression can be derived for the case in which there are only three time periods,  $t \in \{0, 1, 2\}$ . Half the households receive the rebate at  $t = 0$ . They have an average contemporaneous treatment effect of  $\beta^0$  and a lagged treatment effect of  $-\beta^0$ . The other half receive the rebate at  $t = 1$  and have an average contemporaneous treatment effect of  $\beta^1$  and a lagged treatment effect of  $-\beta^1$ . Then one can show that the average and lagged treatment effects are:

$$(25) \quad \beta^{OLS} = \frac{1}{2}(\beta^0 + \beta^1) + \frac{1}{2}(\beta^0 - \beta^1) = \beta^0$$

$$(26) \quad \gamma^{OLS} = -\frac{1}{2}(\beta^0 + \beta^1) + \frac{1}{2}(\beta^0 - \beta^1) = -\beta^1$$

These expressions show that if treatment effects are heterogeneous, then the homogeneous OLS estimator will in general be biased relative to the population treatment effect  $\frac{1}{2}(\beta^0 + \beta^1)$ . The bias will depend on the sign of  $\beta^0 - \beta^1$ , i.e. whether the earlier treatment effects are larger or smaller than the later treatment effects. If  $\beta^0 > \beta^1$ , then there is an upward bias. This is because OLS will use the group 1 smaller reversal at  $t = 2$  to correct for the group 0 larger reversal at  $t = 1$ . This correction is too small since  $-\beta^0 + \beta^1 < 0$ , which implies this counterfactual group will still be contaminated by the lagged treatment effect and inflate the OLS estimates. Furthermore, OLS will also put more weight on earlier treated units which also causes an upward bias if  $\beta^0 > \beta^1$ . In contrast, if  $\beta^0 < \beta^1$  then these two biases reverse and the homogeneous OLS estimator is too small.

To assess the importance of treatment effect heterogeneity in this setting we estimate the following heterogeneous-effects specification:

$$(27) \quad C_{i,t} - C_{i,t-1} = \sum_s \beta_{0,s} month_s + \beta_1' \mathbf{X}_{i,t} + \sum_{e=0}^T \beta_{2,e} I(ESP_{i,t}) I(ESP_{i,e}) \\ + \sum_{e=0}^T \beta_{3,e} I(ESP_{i,t-1}) I(ESP_{i,e}) + u_{i,t}$$

where  $\beta_{2,e}$  is the treatment effect of a cohort that received the rebate at  $t = e$  and  $\beta_{3,e}$  is the corresponding lagged treatment effect. This specification is similar to the solution to heterogeneous treatment effects proposed in Sun and Abraham (2020).

[Table IV Here]

Column (3) of Table III reports estimates of the weighted contemporaneous effect  $\sum_{e=0}^T w_e \beta_{2,e}$  and the weighted lagged effect  $\sum_{e=0}^T w_e \beta_{3,e}$ , where the weights correspond to the OLS weights of the cohorts. Allowing for heterogeneous effects in the full sample reduces the contemporaneous rebate effect by \$90. From Figure VI we know that the early treatment effects in the full sample are larger than the later treatment effects, which causes an upward bias in the contemporaneous effect. By contrast, in the rebate-only sample allowing for heterogeneous effects increases the contemporaneous treatment effect by \$70 because the later treatment effects are larger.

#### IV.D Rebate Reporting Bias

We now discuss a bias associated with the correlation between a household's report of rebate receipt and low spending in the previous period in the CEX sample. To show the correlation, we first regress consumer expenditure on an indicator for receiving a rebate in both the current and the next interview

$$(28) \quad C_{i,t} = \sum_s \delta_{0,s} month_s + \delta_1' \mathbf{X}_{i,t+1} + \delta_2 I(ESP_{i,t}) + \delta_3 I(ESP_{i,t+1}) + u_{i,t}$$

where  $\delta_3$  captures the effect of future rebate receipt on current spending. We estimate this specification in levels to maintain the same sample as our other regressions.

Column (1) of Table IV shows a large negative effect of future rebate receipt on current expenditure. This result likely reflects that rebate recipients have lower average consumption than non-recipients. In column (2) we therefore restrict the estimation to the rebate only sample, in which there should be no such rebate reporting bias. We find that the estimate remains economically very large at -\$575 and statistically significant at the 10% level. This estimate suggests that rebate recipients had unusually low levels of spending in the period before the rebate arrived.

Absent anticipation effects past expenditure should not forecast the timing of rebate if it is randomly assigned.<sup>23</sup> Table IV then implies that the rebate timing in the CEX is not random. Thus, while the true timing of rebates is based on the last two digits of the social security number, which is essentially randomized, the *reported* rebate timing may not be. Consider a household receiving a rebate in May. It should be equally likely sampled by the CEX in either June, July, or August. However, in Appendix Table D.1 we document that households are systematically more likely to report receiving the rebate in the month before the interview (June in this example). We also show in Appendix Table D.2 that households that report receiving the rebate in the month before the interview display a much greater increase in expenditure than households that report receiving the rebate three months ago. This suggests that there could be important recall issues with households more likely to report rebates when they accompany large increases in expenditures, though the exact source is difficult to pin down.

Regardless of the source, the negative correlation between rebate reporting and lagged expenditure will cause an upward bias in the MPC estimates. To illustrate this point, we now distinguish between the true rebate date  $I(ESP_{i,t}^*)$  and the reported rebate date  $I(ESP)_{i,t} = I(ESP_{i,t}^*) + \nu_{it}$  where  $\nu_{it}$  is the reporting error. We also model mean reversion in consumer expenditure, since this is a feature of the CEX and important for the bias:

$$(29) \quad C_{it} = \alpha_i + \lambda_t + \rho C_{i,t-1} + \beta I(ESP_{i,t}^*) + \gamma I(ESP_{i,t-1}^*) + \epsilon_{it},$$

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23. We do not pursue an explanation based on anticipation effects for three reasons: First, we formally test for anticipation effects in appendix section III.F and cannot reject our null hypothesis of no anticipation effects. Second, including the lead of the rebate in our regression specifications has almost no effect on the contemporaneous rebate coefficient in the full sample, while in the presence of negative anticipation effects including the lead of the rebate coefficient should lower the contemporaneous rebate coefficient (e.g., it does so in our Monte Carlo). Third, we are not aware of any model that gives rise to negative anticipation effects.

where  $\rho \in (0, 1)$  determines the degree of mean-reversion. We assume  $\gamma = -\rho\beta$  so the rebate has a one-time effect on the level of consumer expenditures.

If we estimate the regression in changes and omit lagged expenditure from the regression,

$$(30) \quad \begin{aligned} \Delta C_{it} &= \Delta \lambda_t + \beta I(ESP_{i,t}) + \gamma I(ESP_{i,t-1}) + \eta_{it}, \\ \eta_{it} &\equiv \alpha_i + (\rho - 1)C_{i,t-1} + \Delta \epsilon_{it} - \beta v_{it} + \gamma v_{i,t-1}, \end{aligned}$$

then the OLS estimator on the contemporaneous effect is

$$(31) \quad \beta^{OLS} = \beta - \beta \frac{Var(\tilde{v}_{it})}{Var(I(ESP_{it}))} + (\rho - 1) \frac{Cov(\tilde{C}_{i,t-1}, \tilde{v}_{it})}{Var(I(ESP_{it}))} + \rho \frac{Cov(\alpha_i, \overline{I(ESP_{it})})}{Var(I(ESP_{it}))}$$

where  $\tilde{X}_{it}$  is the residual from the regression of the variable  $X_{it}$  on a time fixed effect and  $I(ESP)_{i,t-1}$ . The second term represents classical measurement error in the treatment variable and will bias  $\beta^{OLS}$  towards zero. The third term is a selection bias in reported treatment: Table IV shows that households are more likely to misreport a rebate following a period of low expenditure so  $Cov(\tilde{C}_{i,t-1}, \tilde{v}_{it}) = Cov(\tilde{C}_{i,t-1}, \overline{I(ESP_{it})}) < 0$ . With mean reversion in expenditure  $\rho < 1$ , this will cause an upward bias in  $\beta$  as consumption is expected to increase when the rebate is reported but that increase is due to mean-reversion not the rebate itself. The fourth term represents a selection bias on permanent consumption: If  $\rho \neq 0$ , then first differencing no longer removes the household fixed effect.

To address the selection bias in reported treatment we add lagged consumer expenditure to our regression,

$$(32) \quad \begin{aligned} C_{i,t} - C_{i,t-1} &= \sum_s \beta_{0,s} month_s + \beta'_1 \mathbf{X}_{i,t} + \sum_{e=0}^T \beta_{2,e} I(ESP_{i,t}) I(ESP_{i,e}) \\ &\quad + \sum_{e=0}^T \beta_{3,e} I(ESP_{i,t-1}) I(ESP_{i,e}) + \beta_4 C_{i,t-1} + u_{i,t} \end{aligned}$$

Specifically, we control for both lagged total expenditure and lagged motor vehicle expenditure since we later split spending along these lines and including both controls ensures that our treatment effects add up. We also add controls for income deciles in  $X_{it}$  to mitigate the selection effect on  $\alpha_i$  in the full sample.

Column (4) of Table III shows the implied treatment effects. In the full sample the treatment effect shrinks by \$90 once the lagged control is included. The implied 3-month MPC is 0.28 after we account for all three biases versus 0.5 in the original specification. In the rebate-only sample adding lagged controls shrinks the treatment effect by \$280. The MPC of 0.34 is less than half that in column (1) and very close to our estimates in the full sample. In Appendix D we provide Monte Carlo evidence that various models of selective reporting can quantitatively explain this reduction in our MPC estimates.

Our full sample estimates may still be subject to a selection bias on the household fixed effect—the final term in our OLS bias formula 31. But in the rebate-only sample, that selection bias  $Cov(\alpha_i, \tilde{D}_{it})$  will be zero since all these households are treated at some point in the sample. This suggests that the estimates for the rebate-only sample in column (4) are the most reliable since they account for all the biases we identify.<sup>24</sup> The similarity of the column (4) estimates across both samples suggests that selection on the fixed effect is also unlikely to be important in the full sample. In this sense, column (4) paints a consistent picture that, after correcting for the biases we identify, both estimates round to 0.3.<sup>25</sup>

In Appendix III.G we verify that our preferred specification (32) recovers the true MPCs in household data simulated from the model of section V. In contrast, the estimates from Equation (17) produce upward-biased estimates of the MPC in the simulated household data, consistent with Kaplan and Violante’s (2014) argument that estimates from Equation (17) cannot be interpreted as MPCs.

Finally, in Appendix III.H, we compare the micro-MPC estimates presented here with estimates using the imputation method of Borusyak, Jaravel, and Spiess (2024) (BJS method). An advantage of their imputation method over our OLS specification is that it does not impose a particular dynamic structure of the treatment effects, since it only imposes structure on untreated outcomes. A disadvantage is that the BJS method

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24. Our estimates do not correct for classical measurement error. Our Monte Carlo simulations in Appendix D suggest that this bias is small.

25. Our estimates do not adjust for time aggregation that arises from households receiving the rebate within the three-month period for which they report expenditure. However, our best estimates suggest that time aggregation bias our estimates upwards: First, the six-month MPC is smaller than the three-month MPC. Second, households that receive the rebate two months into the reporting period report much higher expenditure than households that receive the rebate at the beginning of the rebate period. This suggests that an initial positive spending response is followed by negative MPCs, which would bias our three-month estimates upwards. We do not adopt this stance in the model in order to make our counterfactuals more conservative and thus less implausible.

is less efficient than OLS if the model of dynamic treatment effects that we specify in (32) is correct. Ultimately, the BJS method yields micro-MPCs around 0.3, similar to our baseline OLS estimates. This suggests that allowing for more general dynamic treatment effects and heterogeneity than our final estimation equation has only small effects on the implied MPC.

## **IV.E Composition of Spending**

Finally, Table V breaks down the total expenditure response to the rebate into the contribution from motor vehicle spending and other expenditures by estimating equation (32) for each component. The estimates imply that motor vehicle expenditures account for almost all of the total expenditure response. The MPC for motor vehicles is 0.3 in the full sample and 0.26 in the rebate only sample. The importance of vehicle spending is consistent with Adams, Einav, and Levin (2009) who find a substantial increase in car demand during the regular tax rebate season and Aaronson, Agarwal, and French (2012) who document large motor vehicle expenditures following minimum wage hikes. By contrast, we find that there is little change in other expenditures: that MPC is -0.02 in the full sample (column 2) and 0.08 in the rebate only sample (column 4).

[Table V Here]

## **V Revisiting the Model**

We now revisit our general equilibrium model as even the smaller MPCs estimated in the last section suggest implausible general equilibrium counterfactuals when used to calibrate the model from Section III. In this section, we revise that model to incorporate realistic dampening forces. The result is a full reconciliation of the micro MPCs and macro counterfactuals.

Standard New Keynesian models assume that intermediate goods can be frictionlessly turned into either nondurable or durable goods, which implies a constant relative price of 1. Figure VII shows that the relative price of motor vehicles spikes by

around 1.0% in June 2008 after households receive the 2008 rebate relative to January through April 2008.<sup>26</sup> We next show that a model with a realistic calibration of the durable supply elasticity generates similar patterns in the relative price and predicts substantial crowding out of durable expenditure in general equilibrium.<sup>27</sup> And since the last section revealed that virtually all of the spending from the rebate was on motor vehicles, the model also predicts that the aggregate effects of the stimulus are modest.

[ Figure VII Here ]

## V.A The Model with Less Elastic Supply

We generalize the two-good, two-agent New Keynesian model presented in Section III to a model in which the relative supply of durable goods is not perfectly elastic. This part of the model builds on the recent analysis of durable goods expenditures by McKay and Wieland (2021, 2022).

*Durable Goods Production.* Durable goods are produced competitively using non-durables  $N_t$  as inputs,<sup>28</sup>

$$(33) \quad \frac{X_{it}}{p_t^d} = N_{it} \left( \frac{X_t}{\bar{X}} \frac{1}{p_t^d} \right)^{-\zeta}$$

where  $\frac{X_{it}}{p_t^d}$  is the real production of durable goods by firm  $i$  and  $\zeta$  is a negative production externality.  $\zeta$  could alternatively represent a fixed factor of production as in McKay and

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26. We take the flat relative price from January through April as a conservative baseline. An alternative interpretation of Figure VII is that the relative motor vehicle price is on a downward trend. This implies that the upward spike is 1.5% relative to trend. Calibrating the model towards that price change requires a lower supply elasticity of  $\zeta^{-1} = 1$  and implies that the rebate is completely crowded out at the aggregate level. In this sense, our results targeting the level change is conservative.

27. Another possible explanation for the rise in relative prices is a shift in the supply curve caused by strikes from the end of February through late May 2008. The strikes affected production of axles used in General Motor's light trucks. Our reading of Automotive News revealed that a combination of gas-price induced declines in demand for light trucks, ramped up production before the strike to stockpile dealer inventories, and capacity enlargements at a Mexican axle factory resulted in increases in GM's light truck inventories over this period. Thus, the strikes did not appear to constrain supply.

28. In our model all durables are produced domestically. During the rebate period in 2008, 75% of all new vehicles were produced domestically. To the extent that the model does not account for expenditure leakage abroad, it overstates the Keynesian multiplier effect and thus also the transfer multiplier.

Wieland (2021). We model it as a production externality because this yields zero profits in durable production.

Real profits from the sale of durable goods are given by

$$(34) \quad \max_{N_{it}} (X_{it} - N_{it}) = \max_{N_{it}} \left[ p_t^d N_{it} \left( \frac{X_t}{\bar{X}} \frac{1}{p_t^d} \right)^{-\zeta} - N_{it} \right]$$

Profit maximization yields an upward sloping supply curve,

$$(35) \quad p_t^d = \left( \frac{X_t}{\bar{X}} \right)^{\frac{\zeta}{1+\zeta}}$$

where  $\bar{X}$  is steady state durable expenditure, so the steady state relative durable price is normalized to 1. Since durable expenditure is denominated in units of nondurable consumption, the supply elasticity of real durable goods is given by  $\frac{1}{\zeta}$ .<sup>29</sup>

*Calibration.* We calibrate the supply elasticity of durable goods  $\zeta^{-1} = 5$ . This is midway between the elasticities reported in House and Shapiro (2008) and Goolsbee (1998). Their elasticities are for overall investment and are not separately reported for motor vehicles, which is how we interpret durable goods in the model. Nevertheless, setting  $\zeta^{-1} = 5$  allows the model to exactly replicate the 1.0% spike of the relative motor vehicle price in the data. As with the baseline model, we simulate the model for micro MPCs of 0.3, 0.5, and 0.9 and corresponding durable MPCs of 0.3, 0.4, and 0.4.

## **VB General Equilibrium Counterfactuals with Less Elastic Supply**

Figure VIII plots the counterfactuals for the model with less elastic durable supply. The left column is identical to Figure IV because the relative durable price only changes in general equilibrium. In the top right panel, we no longer see evidence of sharp V-shapes for total PCE in the general equilibrium counterfactual for our preferred MPC of 0.3. This counterfactual is spanned by the data and the most pessimistic forecast and is well within the historical norm, so we consider it plausible.

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29. Our model interprets the relative durable price spike in Figure VII as an increase in relative marginal cost. An alternative interpretation of the data is that the elasticity of substitution across cars temporarily fell as a result of the stimulus check causing an increase in mark-ups. This mechanism would also produce equivalent crowding out through higher relative prices. In addition, if motor vehicle profits accrue to optimizing agents, then this mechanism would reduce the covariance between income changes and MPCs in the model and make the stimulus less effective (Auclert, 2019; Bilbie, 2024).

[ Figure VIII Here ]

The two middle panels reveal general equilibrium responses of motor vehicle expenditure to a tax rebate that are much less than the micro MPCs, and this difference largely accounts for the dampening of total PCE in general equilibrium. The general equilibrium dampening of the motor vehicles responses stems from the rise in relative motor vehicle prices. In our preferred calibration with a micro MPC of 0.3, the tax rebate increases the relative vehicle price by 1.0% in July 2008 followed by a gradual decline. The bottom right panel of Figure VIII shows the implied GE counterfactual relative vehicle price absent the rebate. For all three MPC calibrations, the June spike in the counterfactual relative motor vehicle price is muted relative to the data.

Optimizing households intertemporally substitute away from durable goods purchases because their price is temporarily high; however, there is only a small amount of intratemporal substitution toward nondurable goods. Hand-to-mouth households also reduce their real expenditures on durable goods, but in their case, it is because their MPCs are fixed in nominal terms so the rise in relative prices of durable goods eats up part of their spending. Appendix Figure C.8 shows that the non-recipients in the CEX do reduce motor vehicle expenditure exactly when the rebates are given out. This is suggestive evidence for this mechanism, though we cannot rule out that it could have been caused by another shock.

[ Table VI Here ]

Panel A Table VI tabulates the correspondence between the micro MPCs and the GE-MPCs in the model. The first row calibrates to our estimates and the next two rows calibrate to the lower and upper bounds of the Parker, Souleles, Johnson, and McClelland's (2013) estimates. When the micro MPC is 0.3, the total consumption GE-MPC is only 0.06. In this case, the general equilibrium forces of the model dampen the effect of the rebate on consumer expenditure. The table shows, however, that as the micro MPCs rise, the amplifying forces eventually become dominant.

The next four columns of Panel A decompose the MPCs into durable expenditure (motor vehicles) and nondurable expenditure. The decomposition demonstrates that dampening in general equilibrium is concentrated in durable expenditure. When the

micro MPC on durables is 0.3, then the durables expenditure GE-MPC is only 0.09. There is a minor decline in nondurable expenditure, reflecting a small rise in the real interest rate through the central bank's reaction to the rebate.

Panel B of Table VI shows the corresponding micro and GE-MPCs in our baseline model in Section III. Comparing Panels A and B, we find that the lower micro MPC and the lower durable supply elasticity each account for approximately half of the decline in the GE-MPC.<sup>30</sup> We will discuss Panel C in the next section.

Because the durable demand elasticity is an important determinant of crowding out, in appendix Table B.1 we investigate the sensitivity of our results to a less elastic demand based on the lower end of the range of estimated elasticities (e.g. Baker, Kueng, McGranahan, and Melzer (2019)). The GE-MPC is 0.12 when the demand elasticity is -6, as compared to 0.06 in our baseline calibration to an elasticity of -15. Thus, there is still substantial crowding out in general equilibrium.

The combination of these dampening general equilibrium forces and more modest micro MPC estimates yields macroeconomic counterfactuals that we consider plausible, reconciling the implausible counterfactuals based on the original micro estimates and the model in Section III. However, they also imply that the effect of the rebate on consumption expenditures in general equilibrium was modest.

## VI Broader Implications for Micro MPCs and Macro Models

This paper has presented a variety of empirical and theoretical results indicating that the 2008 rebates resulted in minimal aggregate stimulus. In this final section, we first briefly summarize our main findings and then discuss the broader implications for micro MPCs and macro models.

We have argued that the upper bound on the GE-MPC for the 2008 rebates is 0.2, a number based on the cumulative gap between the actual behavior of aggregate consumption and our most pessimistic real-time forecast of its path without the rebates. To

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30. The lower micro MPC interacts nonlinearly with the lower durable supply elasticity, so one cannot uniquely decompose the decline of the GE-MPC. Starting with a micro MPC of 0.5, the lower micro MPC estimates account for 26-44% in the decline of the GE-MPC with the remaining 56-74% due to GE dampening from less elastic durable supply. For a micro MPC of 0.9, the corresponding shares are 65-88% for the lower micro MPC and 12-35% for GE dampening. The midpoint of these ranges is 56% for the lower micro MPC and 44% for GE dampening.

reconcile this small number with micro estimates and macro models, we first corrected several biases in some of the leading micro estimates. Our new estimates indicate a micro MPC from the 2008 rebates of 0.3 for total consumption, with all of it coming from spending on motor vehicles. However, the high micro MPC on motor vehicles did not translate into a boom in aggregate motor vehicle spending. Instead, the combination of an upward sloping relative supply curve with relatively elastic demand for motor vehicle spending resulted in a temporary increase in the relative price of motor vehicles and only a small increase in sales. We estimated a GE-MPC for total consumption of only 0.1 based on our calibrated model. Moreover, since investment and net exports did not change, the GDP multiplier was also about 0.1. Thus, we conclude that Feldstein (2008) and Taylor (2009) were correct: the 2008 tax rebates provided at most a tiny stimulus to the economy.

Our new micro MPC estimates are consistent with estimates from a number of other studies. For example, Sahm, Shapiro, and Slemrod's (2010) Michigan Survey asking what households did with their 2008 rebates implies a total spending MPC of about one-third. Their survey did not distinguish between durable and nondurable goods, but if the MPC came mostly from motor vehicles, then their results would imply a GE-MPC of 0.2 or below. Borusyak, Jaravel, and Spiess's (2024) analysis of the Nielsen data, which corrects two kinds of bias in Broda and Parker's (2014) analysis of those data, produces estimates of quarterly MPCs for the basket of Nielsen products of 3.4 percent rather than Broda and Parker's (2014) 6.7 percent. When they use Broda and Parker's (2014) method to scale up the estimates to all nondurables consumption, they obtain an MPC for all nondurables between 7 and 10 percent rather than Broda and Parker's (2014) 14 to 21 percent. Thus, Borusyak, Jaravel, and Spiess's (2024) small MPC on nondurables is also consistent with a GE-MPC of 0.2 or below. Recently, Boehm, Fize, and Jaravel (forthcoming) conducted an experiment in France involving temporary transfers and estimated a quarterly micro MPC of 0.23, with the bulk of spending on durables or semi-durables.<sup>31</sup> Thus, these results are also consistent with low GE-MPCs.

Some studies, however, find micro MPCs as high as 0.5 for total consumer expenditures. For example, Jappelli and Pistaferri (2014, 2020) estimate an MPC of 0.48 using survey data in Italy that asks households how much they would spend from a hypothetical transitory income change equal to one month's income.<sup>32</sup> Fagereng, Holm,

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31. These results refer to Group 1, whose gift cards had no expiration.

32. The Italian survey has no information on intertemporal MPCs, however. As Jappelli and Pistaferri (2014) write: "the question does not ask consumers over which period they plan to spend or save the

and Natvik (2021) use Norwegian tax registry data to estimate the MPC out of lottery winnings in Norway and estimate an MPC of 0.5 in the year of winning. The Norwegian data do not measure consumption expenditures so the researchers must impute it from income and wealth changes along with assumptions on capital gains.<sup>33</sup> None of these papers distinguishes between nondurables and durable goods. As the second line of Panel A in Table VI shows, a micro MPC for total consumption of 0.5 is consistent with a GE-MPC around 0.2 if 0.4 of that MPC is on durable goods such as motor vehicles. Thus, even these higher estimated MPCs are potentially consistent with our finding of a small GE-MPC as long as most of the induced spending is on motor vehicles and similar durables.

Agarwal and Qian (2014) estimate very high MPCs that are not consistent with small GE-MPCs. They use a 2011 natural experiment in Singapore to estimate cumulative micro MPCs of 0.8 within ten months of the announcement of large cash transfers. In a small open economy such as Singapore, the GE-MPC should be equal to the micro MPC (see Aggarwal, Auclert, Rognlie, and Straub (2023)). Ramey (2025) tests this prediction by constructing a new narrative series and using time series methods that exploit the repeated nature of the natural experiment. She finds that the cash transfers significantly raised aggregate disposable income and saving rates in Singapore, but had no effect on aggregate consumption expenditures. Thus, the GE-MPC estimates were zero despite the high micro MPC estimates.

One study that found sizeable MPCs specifically on nondurables is Johnson, Parker, and Souleles's (2006) analysis of the 2001 rebates. They estimated a quarterly MPC of 0.37 for nondurables but found no significant effects on durable goods or total consumption expenditures in CEX data.<sup>34</sup> In Orchard, Ramey, and Wieland (Forthcoming) we revisit Johnson, Parker, and Souleles's (2006) 2001 rebate estimates using the CEX data. On re-examining the data, we discovered that their definition of nondurables excluded some nondurables (some of which had negative MPCs) but included some

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transfer (i.e., one month, one year, or the entire lifetime)" (p. 125). Indeed, a permanent income consumer with no bequest motive would answer "100%" if they were thinking of their entire lifetime. Interestingly, 100% is the third most frequent answer given, after 50% and 0%.

33. Golosov, Graber, Mogstad, and Novgorodsky (2024) use IRS data to estimate an annual MPC of 0.6 out of lottery winnings in the U.S. However, their imputed measure of consumption includes paying down debt, which numerous studies (e.g. Sahm, Shapiro, and Slemrod (2010) and Hisnanick and Kern (2018)) have found to be an important use of transitory increases in income.

34. A number of papers including Hamilton (2008), Misra and Surico (2014), and Kaplan and Violante (2014) trim extreme values of consumption growth in the sample to lower the nondurables MPC estimate towards 0.25.

durables. When we instead use the BLS' correspondence between CEX expenditure categories and BEA consumer expenditure definitions, we estimate MPCs near zero on nondurables as well as total consumption. Thus, we find no evidence of stimulus from the 2001 rebates.

As the previous discussion highlights, our new model with less elastic motor vehicle supply reveals that the aggregate effects of a temporary rebate depend importantly on the distribution of spending across nondurable and durable goods, and not only on the average MPC. To illustrate this point, Panel C of Table VI shows the GE-MPC in a model that abstracts from durable goods and calibrates the nondurable micro MPC to the overall response to expenditure.<sup>35</sup> In this model, when the micro MPC for nondurable expenditure (and thus overall expenditure) is 0.30, the GE-MPC is 0.36. Thus abstracting from durable goods yields the conclusion that a tax rebate is amplified in general equilibrium. By contrast, in our model with durable goods the GE-MPC is only 0.07 (Panel A), which is significantly smaller than the GE-MPC in the nondurables-only model. To obtain this GE-MPC in a nondurables-only model the micro MPC must be calibrated to 0.07, much smaller than the conventional 0.25, which would generate a GE-MPC close to 0.3. This sizeable difference stems from the combination of upward-sloping relative supply of durables with durable demand being more elastic than nondurable demand.

Thus, the composition of consumer spending is an important determinant of the aggregate effects of transfers. Auclert, Rognlie, and Straub (2024), Laibson, Maxted, and Moll (2022), and Wolf (2023) incorporate durables in versions of their HANK models, but they assume perfect substitutability in production and hence their models feature a constant relative price of durables. Thus, their model misses the important general equilibrium dampening effect that durable goods can create. In contrast, our results parallel Boehm's (2020) finding that short-run government consumption multipliers are greater than government investment multipliers. In his model, government investment raises the relative price of investment goods, which crowds out private investment because of its high intertemporal elasticity of substitution.

A remaining question is the dynamic path of micro MPCs. Our estimates and those of Sahm, Shapiro, and Slemrod (2010), Borusyak, Jaravel, and Spiess (2024), and others imply that micro MPCs for the 2008 rebate are positive for at most a few months. Similarly, recently work by Boehm, Fize, and Jaravel (forthcoming), who conduct a ran-

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35. In this model we set the weight on the utility of durables stock to  $\psi = 0$ , the durable operating cost to  $\eta = 0$ , and the MPC for durables to  $mpx = 0$ .

domized control experiment of temporary transfers to French households, finds that all the spending response is concentrated in the first three weeks. In contrast, Fagereng, Holm, and Natvik (2021) estimate positive MPCs that persist for several years for lottery winners in Norway and Colarieti, Mei, and Stantcheva (2024) elicit positive intertemporal MPCs for at least a year in a survey of U.S. households. The Norwegian estimates form the basis for calibrating intertemporal MPCs in Auclert, Rognlie, and Straub's (2024) HANK model. Understanding why the intertemporal MPCs vary across these studies is an important avenue for future research.

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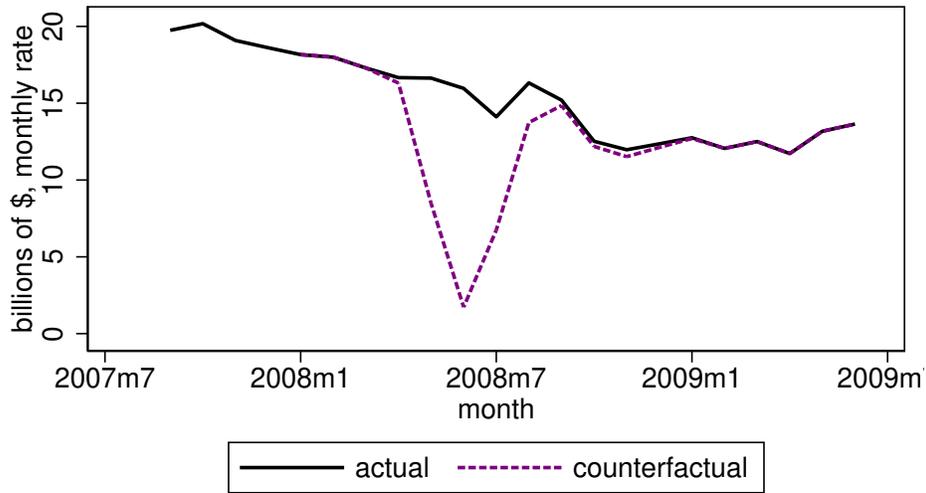
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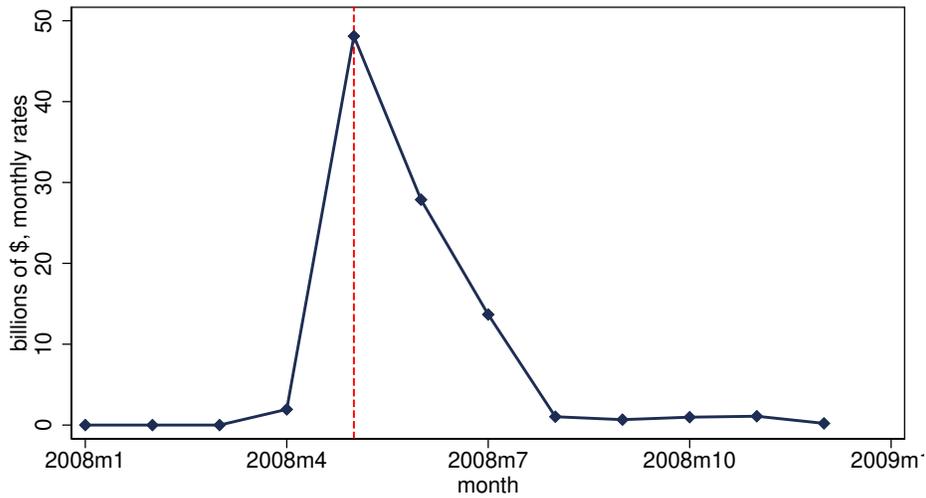
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## Figures



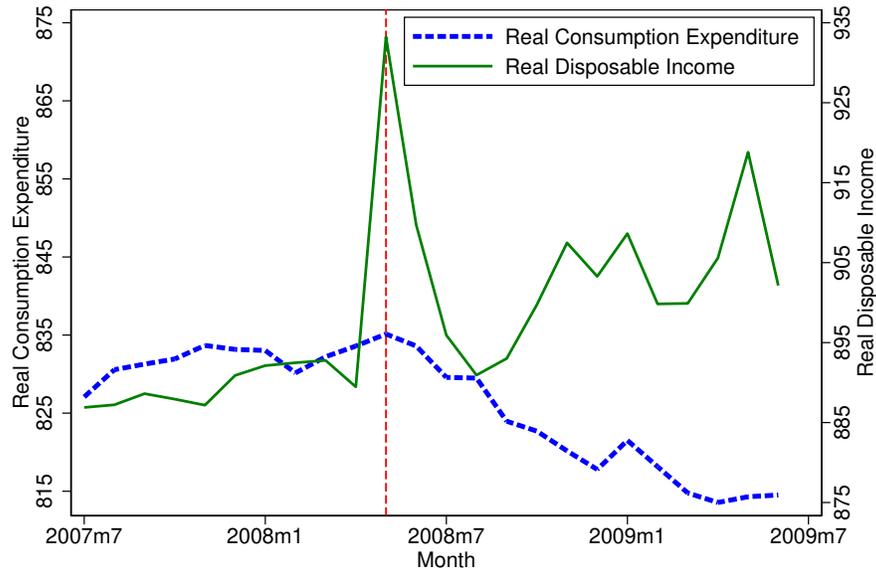
**Figure I**  
**Expenditures on New Motor Vehicles: Actual vs. Counterfactual**

Based on Sahm, Shapiro, and Slemrod (2012) calculations applied to revised data.



**Figure II**  
**2008 Tax Rebates**

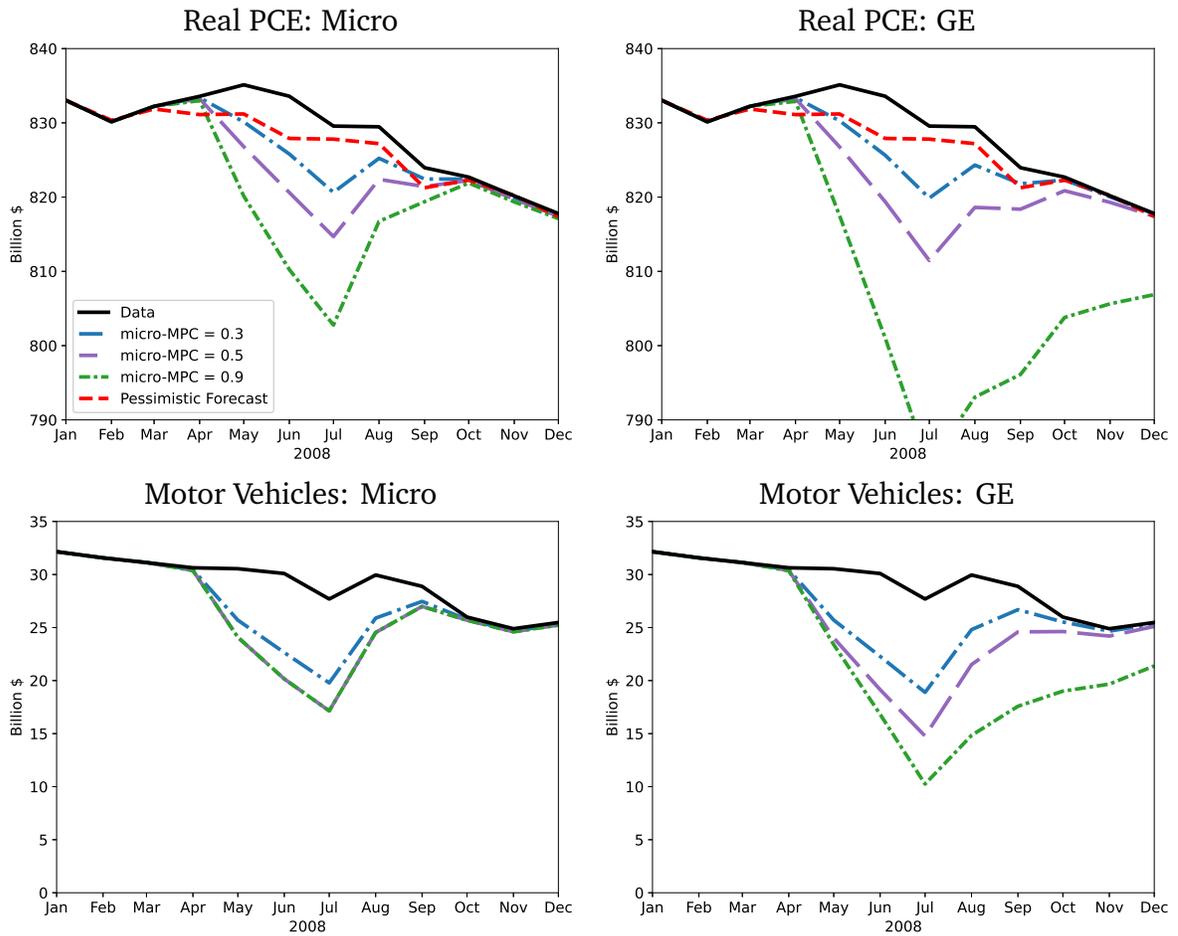
Data from Shapiro and Slemrod (2009). Rebates are nominal. Vertical dashed line indicates May 2008.



**Figure III**

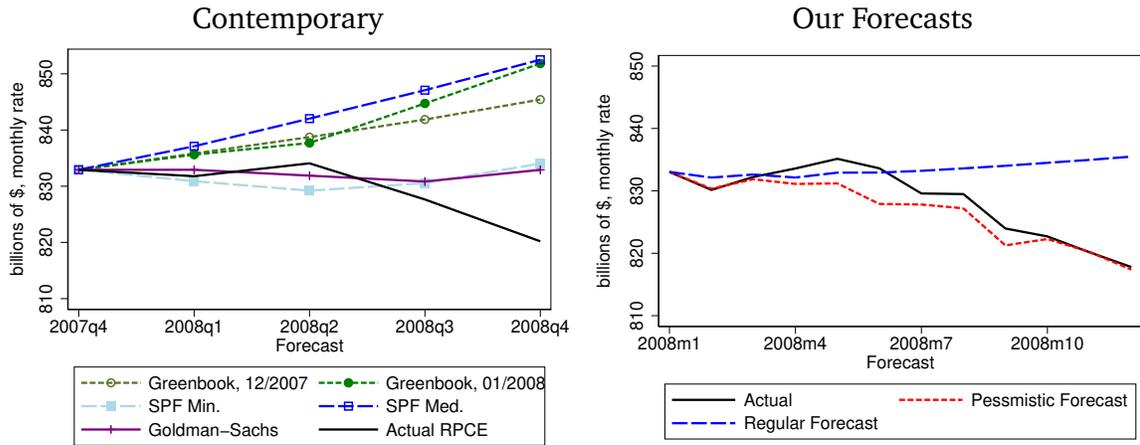
**Real Aggregate Disposable Income and Consumption Expenditure**

Source. BEA data. Vertical dashed line indicates May 2008.



**Figure IV**  
**Counterfactual Real Consumption Expenditures: Baseline Model**

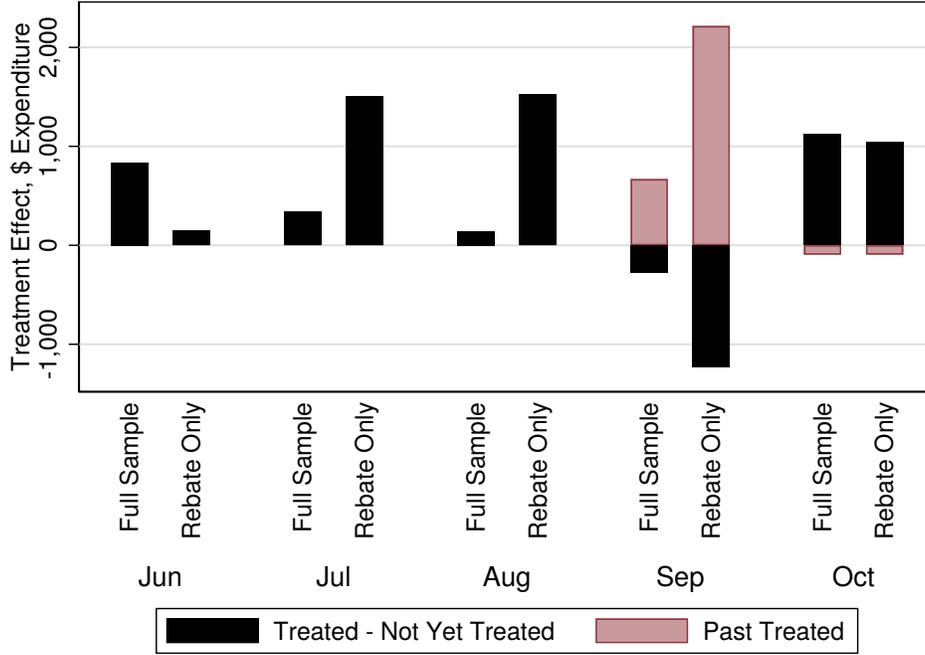
Based on Two-Good, Two-Agent NK model simulations and actual data on rebates and consumption. The micro MPC value refers to the MPC for total consumption.



**Figure V**  
**Real Consumption Forecasts**

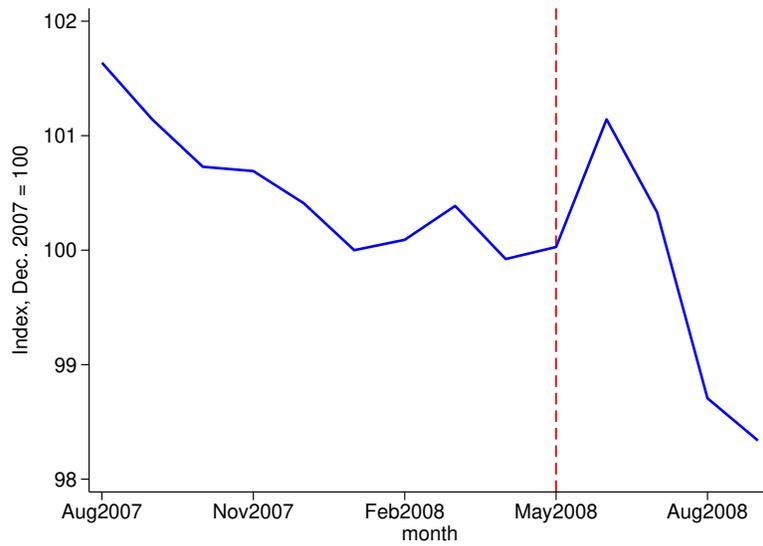
Left Panel: Sourced from BEA data, Federal Reserve Board, Survey of Professional Forecasters. All forecasts normalized to monthly real consumption in 2007Q4.

Right Panel: Forecasts are based on information through January 2008, with exception of models in which oil prices are assumed to follow their actual path and Lehman Brothers dummies are included. Real oil prices exogenously follow their actual path in the pessimistic forecast and their forecasted path in the regular forecast; recession dummies and Lehman Brothers bankruptcy dummy variables are included only in the pessimistic forecast.



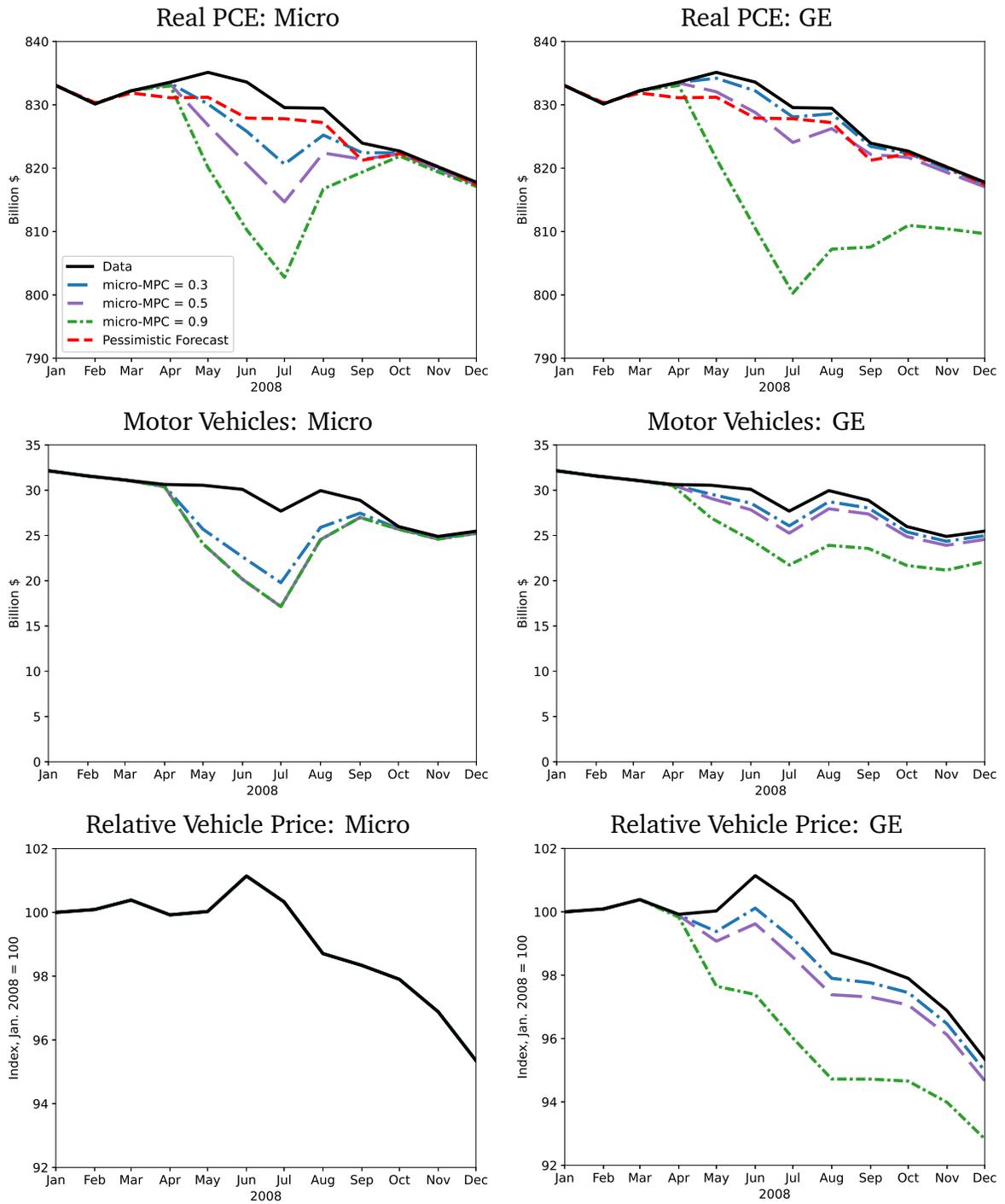
**Figure VI**  
**TWFE Coefficients in the Full and Rebate Only Samples By Month**

The dependent variable is the change in PCE. Periods after October, 2008, also receive positive weight, however, these weights small and not shown here.



**Figure VII**  
**Motor Vehicle Relative Prices**

Source: BLS research CPI for new motor vehicles divided by core CPI (Williams and Sager, 2019).



**Figure VIII**

**Counterfactual Real Consumption Expenditures: Less Elastic Supply**

Based on NK model simulations and actual data on rebates and consumption. The micro MPC value refers to the MPC for total consumption.

## Tables

**Table I**  
**Baseline Calibration of the Model**

Parameter	Value	Description
$\beta$	0.997	Subjective discount factor
$\psi$	0.175	Weight on durable service flow
$\sigma$	0.5	Utility curvature on nondurable consumption
$\sigma^d$	1	Utility curvature on durable service flow
$\theta^d$	varies	Calvo parameter on durable adjustment
$\eta$	0.018	Durable operating cost
$\nu$	76.918	Weight on disutility of labor
$\phi$	1	Inverse of the Frisch elasticity of labor supply
$\gamma$	varies	Fraction of Hand-to-Mouth consumers
$\vartheta$	varies	Hand-to-Mouth MPC on durables
$\delta^d$	0.015	Depreciation of durable consumption goods
$\alpha$	0.36	Exponent on private capital in production function
$\delta$	0.005	Depreciation of private capital
$\kappa$	40	Investment adjustment cost parameter
$\delta_1$	0.008	Parameter on linear term of capital utilization cost
$\delta_2$	0.017	Parameter on quadratic term of capital utilization cost
$\mu_p$	1.2	Steady-state price markup
$\mu_w$	1.2	Steady-state wage markup
$\theta_p$	0.917	Calvo parameter on price adjustment
$\theta_w$	0.917	Calvo parameter on wage adjustment
$\epsilon_p$	6.0	Elasticity of substitution between types of goods
$\epsilon_w$	6.0	Elasticity of substitution between types of labor
$g_Y$	0.2	Steady-state share of total govt spending to GDP
$\phi_b$	0.027	Debt feedback coefficient in fiscal rule
$\rho_r$	0.947	Monetary policy interest rate smoothing
$\phi_\pi$	1.5	Monetary policy response to inflation
$\phi_{gap}$	0.083	Monetary policy response to the output gap

*Notes:* The model is calibrated at a monthly frequency. The parameter  $\gamma$  is calibrated to either 0.3, 0.5, or 0.9, which corresponds to the micro MPC in the model. The parameter  $\theta^d$  is calibrated such that for each value of  $\gamma$  to model replicates our empirical targets for the short-term interest elasticity of durable demand. For example, when  $\gamma = 0.3$ , then  $\theta^d = 0.85$ . The parameter  $\vartheta$  is calibrated to match an overall MPC on motor vehicles of 0.3 when  $\gamma = 0.3$  and of 0.4 when  $\gamma = 0.5$  or  $\gamma = 0.9$ . This yields  $\vartheta = 1.0$  when  $\gamma = 0.3$ ,  $\vartheta = 0.8$  when  $\gamma = 0.5$ , and  $\vartheta = 0.44$  when  $\gamma = 0.9$ . See the text for details.

**Table II**  
**Model Counterfactuals Compared to Largest Historical Expenditure Decline**

<b>Panel A: Total PCE</b>				
Largest Historical Declines			Model Counterfactuals	
Date	Episode	Decline	Calibration	Decline
Jan-Apr 2020	COVID lockdowns	16.9	micro MPC = 0.9	5.9
Jan-Apr 1980	Credit controls, Volcker	2.9	micro MPC = 0.5	2.6
Aug-Nov 1974	prior spike up	2.3	micro MPC = 0.3	1.6
Apr-Jul 1960	prior spike up	1.8		
Sep-Nov 2008	Lehman Collapse	1.1		

<b>Panel B: Motor Vehicle Expenditures</b>				
Largest Historical Declines			Model Counterfactuals	
Date	Episode	Decline	Calibration	Decline
Jan-Apr 2020	COVID lockdowns	31.4	micro MPC = 0.9	66.2
Aug-Nov 1974	prior spike up	25.3	micro MPC = 0.5	51.4
Jul-Oct 2005	prior spike up	25.3	micro MPC = 0.3	38.0
Jan-Apr 1980	Credit controls, Volcker	24.8		
Sep-Nov 2008	Lehman Collapse	16.9		

Top Panel: Four largest three-month declines of personal consumption expenditures (left three columns) compared to three-month decline implied by model counterfactual from April through July 2008 for a total micro MPC of 0.3, 0.5 or 0.9 (right two columns). The Lehman collapse is added as an additional comparison.

Bottom Panel: Four largest three-month declines of motor vehicle expenditures (left three columns) compared to three-month decline implied by model counterfactual from April through July 2008 for a total micro MPC of 0.3, 0.5 or 0.9 (right two columns). The Lehman collapse is added as an additional comparison.

**Table III**  
**Household PCE Response to Rebate**

<b>Panel A: Full Sample</b>				
	Homogeneous Treatment		Heterogeneous Treatment	
	(1)	(2)	(3)	(4)
Rebate Indicator	470.13** (213.56)	433.84** (206.72)	347.40* (211.03)	261.99 (182.84)
Lag Rebate Indicator		-173.61 (222.27)	-190.13 (224.06)	-131.32 (180.89)
Lag Total Expenditure				-0.26*** (0.03)
Lag Motor Vehicle				-0.74*** (0.03)
Implied 3-month MPC	0.50	0.46	0.37	0.28
Implied 6-month MPC		0.72	0.53	0.12
6-Month MPC S.E.		(0.48)	(0.49)	(0.38)
Income Decile FE	No	No	No	Yes
Observations	16,962	16,962	16,962	16,962
<b>Panel B: Rebate Recipients Only</b>				
	Homogeneous Treatment		Heterogeneous Treatment	
	(1)	(2)	(3)	(4)
Rebate Indicator	764.46** (314.81)	527.34 (342.40)	601.27 (403.52)	320.72 (496.73)
Lag Rebate Indicator		-428.09 (354.06)	-478.71 (380.50)	-512.83 (389.05)
Lag Total Expenditure				-0.29*** (0.02)
Lag Motor Vehicle				-0.71*** (0.03)
Implied 3-month MPC	0.82	0.56	0.64	0.34
Implied 6-month MPC		0.65	0.77	-0.15
6-Month MPC S.E.		(0.92)	(1.08)	(1.19)
Income Decile FE	No	No	No	Yes
Observations	10,076	10,076	10,076	10,076

*Notes:* The dependent variable is the change in Personal Consumption Expenditure (PCE). Regressions include interview (time) fixed effects, and household level controls for age, change in number of adults, and change in number of children. Standard errors for the 6-month MPC are estimated via Delta-method. The rebate coefficients in columns (3) and (4) are the weighted average of the interaction between rebate cohort and the (lagged) rebate indicator with weights computed following Sun and Abraham (2021). Standard errors, in parentheses, are clustered at the household level: \*  $p < 0.1$ , \*\*  $p < 0.05$ , \*\*\*  $p < 0.01$ .

**Table IV**  
**Negative effect of future rebate receipt on current expenditure**

	Full Sample (1)	Rebate Recipients Only (2)
Lead Rebate Indicator	-863.0*** (289.0)	-574.9* (332.5)
Rebate Indicator	-414.5 (299.2)	200.5 (368.4)
Observations	16,962	10,076

*Notes:* The dependent variable is the Level of PCE. Regressions include interview (time) fixed effects, and household level controls for age, change in number of adults, and change in number of children. Standard errors, in parentheses, are clustered at the household level: \*  $p < 0.1$ , \*\*  $p < 0.05$ , \*\*\*  $p < 0.01$ .

**Table V**  
**Household Spending Response to Rebate by Subcategory:**

	Full Sample		Rebate Only Sample	
	Motor Vehicles (1)	Other PCE (2)	Motor Vehicles (3)	Other PCE (4)
Rebate Indicator	283.64*** (107.33)	-21.65 (145.54)	249.55 (165.03)	71.17 (459.45)
Lag Rebate Indicator	125.47 (103.53)	-256.80* (141.97)	104.85 (157.92)	-617.68* (355.27)
Lag Total Expenditure	0.02*** (0.01)	-0.28*** (0.03)	0.02*** (0.01)	-0.32*** (0.02)
Lag Motor Vehicle	-1.04*** (0.01)	0.30*** (0.03)	-1.04*** (0.01)	0.33*** (0.03)
Implied 3-month MPC	0.30	-0.02	0.26	0.08
Income Decile FE	Yes	Yes	Yes	Yes
Observations	16,962	16,962	10,076	10,076

*Notes:* Standard errors, in parentheses, are clustered at the household level. Significance is indicated by: \*  $p < 0.1$ , \*\*  $p < 0.05$ , \*\*\*  $p < 0.01$ . All regressions include interview (time) fixed effects, as well as household level controls for age, change in number of adults, and change in number of children. Rebate sample includes only households that receive a rebate at some point during our sample period. The rebate coefficients are the weighted average of the interaction between the rebate cohort and a (lagged) rebate indicator where the weights are derived from Sun and Abraham (2021).

**Table VI**  
**General Equilibrium Marginal Propensity to Consume:**

**Panel A: Model with Less Elastic Durable Supply**

PCE		Motor vehicles		Nondurable goods	
micro	GE	micro	GE	micro	GE
0.30	0.06	0.30	0.09	0.00	-0.03
0.50	0.22	0.40	0.15	0.10	0.07
0.90	1.42	0.40	0.45	0.50	0.97

**Panel B: Baseline Model**

PCE		Motor vehicles		Nondurable goods	
micro	GE	micro	GE	micro	GE
0.30	0.32	0.30	0.34	0.00	-0.02
0.50	0.64	0.40	0.52	0.10	0.12
0.90	2.16	0.40	0.94	0.50	1.22

**Panel C: Model without Durable Goods**

PCE		Motor vehicles		Nondurable goods	
micro	GE	micro	GE	micro	GE
0.06	0.06	0	0	0.06	0.06
0.30	0.36	0	0	0.30	0.36
0.50	0.75	0	0	0.50	0.75
0.90	2.54	0	0	0.90	2.54

MICRO MPCs AND MACRO COUNTERFACTUALS:  
THE CASE OF THE 2008 REBATES—Online Appendix\*

Jacob D. Orchard  
Valerie A. Ramey  
Johannes F. Wieland

## A Model

### I.A Optimizing Households

A measure  $1 - \gamma$  of ex-ante identical households maximizes utility subject to their budget constraints. The utility function of each household  $i$  is

$$E_0 \sum_{t=0}^{\infty} \beta^t \left[ \frac{C_t^o(i)^{1-\frac{1}{\sigma}}}{1-\frac{1}{\sigma}} + \psi \frac{D_t^o(i)^{1-\frac{1}{\sigma^d}}}{1-\frac{1}{\sigma^d}} - \nu \frac{H_t^o(i)^{1+\phi}}{1+\phi} \right]$$

where  $C_t^o(i)$  is nondurable consumption,  $D_t^o(i)$  is the durable stock, and  $H_t^o(i)$  is hours worked.

We assume that optimizing households face an adjustment friction on durable goods, since otherwise they would exhibit extremely high willingness to intertemporally substitute durables purchases. While households optimize their nondurable consumption every period, they do not optimize their durable holdings every period because they face an Evans and Ramey (1992) type of calculation cost. In particular, individual households experience random variations in the psychic costs of calculating optimal durable goods stocks, which could be due to varying cognitive demands of other events in their daily lives, etc. Only a fraction  $1 - \theta^d$  draw costs that are low enough to allow them

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to calculate and hence reoptimize their current durable stock. This friction produces a reversal in durable spending consistent with the evidence (McKay and Wieland (2021)) and keeps the model tractable since it produces a Calvo-type reduced form.

The friction on durable purchases implies that households will generally hold different durables stocks,  $D_t^o(i) \neq D_t^o(j)$ . We assume optimizing households form a family that provides consumption insurance across household members so nondurable consumption is identical,  $C_t^o(i) = C_t^o$ ,  $\forall i$ . Labor supply is not chosen by the household, but instead by a union as discussed below. The union sets labor supply to be equal across households so  $H_t^o(i) = H_t^o$ ,  $\forall i$ .

Integrating across all optimizing households, the utility function for the family is:

$$E_0 \sum_{t=0}^{\infty} \beta^t \left[ \frac{(C_t^o)^{1-\frac{1}{\sigma}}}{1-\frac{1}{\sigma}} + \psi \frac{\int_0^1 D_t^o(i)^{1-\frac{1}{\sigma^d}} di}{1-\frac{1}{\sigma^d}} - \nu \frac{(H_t^o)^{1+\phi}}{1+\phi} \right]$$

The aggregate household budget constraint is

$$\begin{aligned} A_t^o &= \frac{R_{t-1}}{\Pi_t} A_{t-1}^o - C_t^o + W_t H_t^o - X_t^o - OC_t^o - T_t^o + \text{Profits}_t^k + \text{Profits}_t^s \\ X_t^o &= p_t^d \left[ \int_0^1 [D_t^o(i) - (1 - \delta^d) D_{t-1}^o(i)] di \right] \\ OC_t^o &= \eta \int_0^1 D_t^o(i) di \end{aligned}$$

where  $R_t$  is the gross nominal interest rate,  $\Pi_t$  is the gross inflation rate measured in nondurable goods prices,  $A_t^o$  are holdings of the nominal bond,  $W_t$  is the real wage,  $T_t^o$  are net taxes (i.e. taxes less transfers),  $\text{Profits}_t^k$  are profits of the capital good producing firms, and  $\text{Profits}_t^s$  are profits of the sticky-price firms, which produce nondurable goods.  $X_t^o$  is net durable expenditure denominated in nondurable goods, which is the sum of net durable purchases of each household,  $D_t^o(i) - (1 - \delta^d) D_{t-1}^o(i)$ .  $OC_t^o$  are operating costs for the durable durable good (e.g., gasoline) which is a fraction  $\eta$  of the total durable stock held by all households. The inclusion of operating expenditures helps produce more realistic elasticities of durable demand.

The family picks an optimal plan  $\{C_t^o, A_t^o, D_t^o(i)\}_{t=0}^{\infty}$  to maximize utility. The first order conditions for nondurable consumption and assets are:

$$\begin{aligned} \lambda_t &= (C_t^o)^{-\frac{1}{\sigma}} \\ \lambda_t &= \beta \frac{R_t}{\Pi_{t+1}} \lambda_{t+1} \end{aligned}$$

where  $\lambda$  is the Lagrange multiplier on the household budget constraint.

The details of the Calvo adjustment frictions are analogous to those in price or wage setting. We first derive the optimal choice of  $D_t^o(i)$  conditional on being able to adjust. Because the durable stock of household  $i$  in the problem is separable from the durable stock of other households, the optimization problem for household  $i$  is simply,

$$\max_{D_t(i)} \sum_{s=0}^{\infty} (\beta\theta^d)^s \left[ \psi \frac{[(1-\delta^d)^s D_t(i)]^{\frac{\sigma^d-1}{\sigma^d}}}{1-\frac{1}{\sigma^d}} - \lambda_{t+s} \eta (1-\delta^d)^s D_t(i) \right] - \lambda_t p_t^d D_t(i) + \sum_{s=1}^{\infty} \beta^s (\theta^d)^{s-1} (1-\theta^d) \lambda_{t+s} p_{t+s}^d (1-\delta^d)^s D_t(i)$$

Here  $(\theta^d)^s$  is the survival probability of the current durable stock into period  $s$ ,  $\psi \frac{D_t(i)^{\frac{\sigma^d-1}{\sigma^d}}}{1-\frac{1}{\sigma^d}}$  is its contribution to household utility,  $\lambda_{t+s} \eta (1-\delta^d)^s D_t(i)$  is the operating cost while the durable stock remains in place measured in utils,  $\lambda_t p_t^d D_t(i)$  is the purchasing price in utils, and  $\lambda_{t+s} p_{t+s}^d (1-\delta^d)^s D_t(i)$  is the resale value of the durable in utils if another adjustment opportunity arises at time  $t+s$ .

The first order condition for  $D_t(i)$  is then

$$\psi \sum_{s=0}^{\infty} [\beta\theta^d (1-\delta^d)^{\frac{\sigma^d-1}{\sigma^d}}]^s D_t(i)^{-\frac{1}{\sigma^d}} = p_t^d \lambda_t + \eta \sum_{s=0}^{\infty} [\beta\theta^d (1-\delta^d)]^s \lambda_{t+s} - \beta(1-\theta^d)(1-\delta^d) \sum_{s=1}^{\infty} [\beta\theta^d (1-\delta^d)]^{s-1} p_{t+s}^d \lambda_{t+s}$$

The problem is identical across households that can make an adjustment at time  $t$ . Therefore, let  $D_t^{o*}$  denote the common optimal reset value for the durable stock at time  $t$ . The optimal reset value is:

$$D_t^{o*} = \left( \frac{\psi \sum_{s=0}^{\infty} [\beta\theta^d (1-\delta^d)^{\frac{\sigma^d-1}{\sigma^d}}]^s}{p_t^d \lambda_t + \eta \sum_{s=0}^{\infty} [\beta\theta^d (1-\delta^d)]^s \lambda_{t+s} - \beta(1-\theta^d)(1-\delta^d) \sum_{s=1}^{\infty} [\beta\theta^d (1-\delta^d)]^{s-1} p_{t+s}^d \lambda_{t+s}} \right)^{\sigma^d}$$

The first order condition for  $D_t^{o*}$  can be written recursively as,

$$D_t^{o*} = \left( \frac{\Omega_{1t}}{\Omega_{2t}} \right)^{\sigma^d}$$

$$\Omega_{1t} = \psi + \beta\theta^d(1 - \delta^d)^{\frac{\sigma^d - 1}{\sigma^d}} \Omega_{1,t+1}$$

$$\Omega_{2t} = (p_t^d + \eta)\lambda_t - \beta(1 - \delta^d)p_{t+1}^d \lambda_{t+1} + \beta\theta^d(1 - \delta^d)\Omega_{2,t+1}$$

$$= \lambda_t \left[ p_t^d + \eta - \frac{(1 - \delta^d)\Pi_{t+1}}{R_t} p_{t+1}^d \right] + \beta\theta^d(1 - \delta^d)\Omega_{2,t+1}$$

where  $\Omega_1$  is the expected present discounted value of a unit of durable varieties and  $\Omega_2$  is the expected present discounted value of the user cost.

By defining the total durable stock among optimizing households as

$$D_t^o \equiv \int_0^1 D_t^o(i) di,$$

we obtain the standard durable accumulation equation and durable net expenditure as a function of aggregate variables only,

$$D_t^o = (1 - \delta^d)D_{t-1}^o + \frac{X_t^o}{p_t^d}$$

$$X_t^o = p_t^d(1 - \theta^d)[D_t^{o*} - (1 - \delta^d)D_{t-1}^o].$$

Using a log-linear approximation to the first order conditions, the elasticity of durable expenditure with respect to the real interest rate is

$$\frac{d \ln X^o}{d \ln R} = \sigma^d \left[ \frac{1 - \theta^d(1 - \delta^d)}{\delta^d} \right] \left[ \frac{(1 - \delta^d)[1 - \beta\theta(1 - \delta)]}{R - 1 + \delta + \eta} \right]$$

The first term in brackets captures the extensive margin response: When  $\theta^d > 0$  only a fraction of the durable stock can respond to changes in the real interest rate. The second term in brackets captures the intensive margin. Because of the Calvo friction, households know that any durable purchase cannot be immediately sold next period. Therefore the expected user cost of a durable purchase is not just the contemporaneous user cost  $p_t^d + \eta - \frac{(1 - \delta^d)p_{t+1}^d \Pi_{t+1}}{R_t}$  but the whole expected present discounted value  $\Omega_{2t}$ . A short-term change in the real rate has a smaller effect on the expected present discounted value  $\Omega_{2t}$  because the contemporaneous user cost only accounts for a part of it. Therefore the intensive margin also becomes less sensitive to short-term changes in interest rates because these have a smaller effect on the expected user cost of the durable.

## I.B Hand-to-Mouth Households

In order for lump-sum transfers to have general equilibrium effects, we require non-Ricardian households. We adopt Galí, López-Salido, and Vallés's (2007) assumption that a certain fraction  $\gamma$  consume hand-to-mouth. Relative to their set-up, our hand-to-mouth households may consume their income over several periods rather than all at once.

We assume that in steady state, hand-to-mouth households have the same after-tax income and consume the same relative amount of durable and nondurable services as optimizing households,

$$\begin{aligned} WH^m - T^m &= WH^o + \text{Profits}^k + \text{Profits}^s - T^o \\ \frac{C^m}{X^m} &= \frac{C^o}{X^o} \end{aligned}$$

where variables superscripted by  $m$  denote the hand-to-mouth household.

We then directly specify dynamic marginal propensities to consume for nondurable and durable expenditures to match both the allocation across goods and any lagged effects implied by the micro MPC estimates,

$$\begin{aligned} C_t^m - C^m + \eta(D_t^m - D^m) &= \sum_{l=0}^L mpc_l [W_{t-l}H_{t-l}^m - T_{t-l}^m - (WH^m - T^m)] \prod_{k=1}^l \frac{R_{t-k}}{\Pi_{t-k+1}} \\ X_t^m - X^m &= \sum_{l=0}^L mpx_l [W_{t-l}H_{t-l}^m - T_{t-l}^m - (WH^m - T^m)] \prod_{k=1}^l \frac{R_{t-k}}{\Pi_{t-k+1}} \\ 1 &= \sum_{l=0}^L (mpc_l + mpx_l) \\ mpx_l &= \frac{\vartheta}{1 - \vartheta} mpc_l, \quad \forall l = 0, \dots, L \end{aligned}$$

where  $mpc_l$  is the marginal propensity to spend on nondurable goods today out of income  $l$  periods ago, and  $mpx_l$  is the marginal propensity to spend on durable goods today out of income  $l$  periods ago. Income that was saved  $l$  periods ago for consumption today accrues real interest  $\prod_{k=1}^l \frac{R_{t-k}}{\Pi_{t-k+1}}$ .

The marginal utility to consume for the hand-to-mouth household is

$$\lambda_t^m = (C_t^m)^{-\frac{1}{\sigma}}$$

The durable stock owned by the hand-to-mouth consumers follow an analogous accumulation equation

$$D_t^m = (1 - \delta^d)D_{t-1}^m + \frac{X_t^m}{P_t^d}$$

## I.C Wages

A continuum of unions indexed by  $j$  provide differentiated labor services to the final good firm that are substitutable with elasticity  $\epsilon^w$ . Each period there is a iid probability  $\theta^w$  that the union cannot adjust the contract wage. In this case, wages will adjust by a fraction  $\chi^w$  of last periods inflation.

The union imposes the same work hours on optimizing and hand-to-mouth households:

$$H_t^m = H_t^o = H_t$$

The demand for hours from union  $j$  at time  $t + s$  conditional on having last reset wages at time  $t$  is

$$H_{t+s}^d(j) = H_{t+s}^d \left( \frac{W_t(j) \left( \frac{P_{t+s-1}}{P_{t-1}} \right)^{\chi^w} \left( \frac{P_t}{P_{t+s}} \right)^{-\epsilon^w}}{W_{t+s}} \right)^{-\epsilon^w} = H_{t+s}^d W_{t+s}^{\epsilon^w} \left( \frac{P_{t+s}}{P_t} \right)^{\epsilon^w} \left( \frac{P_{t+s-1}}{P_{t-1}} \right)^{-\epsilon^w \chi^w} W_t(j)^{-\epsilon^w}$$

where  $P_t$  is the price level at time  $t$ .

If the union can adjust its wage at time  $t$  it picks the optimal wage to maximize the expected discounted utility of the representative household while this wage prevails:

$$\max_{W_t^*} \sum_{s=0}^{\infty} (\beta \theta^w)^s H_{t+s}^d W_{t+s}^{\epsilon^w} \left( \frac{P_{t+s}}{P_t} \right)^{\epsilon^w} \left( \frac{P_{t+s-1}}{P_{t-1}} \right)^{-\epsilon^w \chi^w} \left[ \tilde{\lambda}_{t+s} \left( \frac{P_{t+s-1}}{P_{t-1}} \right)^{\chi^w} \left( \frac{P_{t+s}}{P_t} \right)^{-1} (W_t^*)^{1-\epsilon^w} - \nu H_{t+s}^\phi (W_t^*)^{-\epsilon^w} \right]$$

where  $\tilde{\lambda} = (1 - \gamma)\lambda_t + \gamma\lambda_t^m$

The first order condition for the union is:

$$\begin{aligned} & (\epsilon^w - 1) \sum_{s=0}^{\infty} (\beta \theta^w)^s H_{t+s}^d W_{t+s}^{\epsilon^w} \left( \frac{P_{t+s}}{P_t} \right)^{\epsilon^w - 1} \left( \frac{P_{t+s-1}}{P_{t-1}} \right)^{-\chi^w (\epsilon^w - 1)} \tilde{\lambda}_{t+s} (W_t^*)^{1-\epsilon^w} \\ & = \epsilon^w \nu \sum_{s=0}^{\infty} (\beta \theta^w)^s H_{t+s}^d H_{t+s}^\phi W_{t+s}^{\epsilon^w} \left( \frac{P_{t+s}}{P_t} \right)^{\epsilon^w} \left( \frac{P_{t+s-1}}{P_{t-1}} \right)^{-\epsilon^w \chi^w} (W_t^*)^{-\epsilon^w} \end{aligned}$$

We write it recursively using

$$\begin{aligned}
F_{1t} &= \nu H_t^d H_t^\phi W_t^{\epsilon^w} (W_t^*)^{-\epsilon^w} + \beta \theta^w \Pi_{t+1}^{\epsilon^w} \Pi_t^{-\chi^w \epsilon^w} \left( \frac{W_t^*}{W_{t+1}^*} \right)^{-\epsilon^w} F_{1,t+1} \\
F_{2t} &= H_t^d W_t^{\epsilon^w} \tilde{\lambda}_t (W_t^*)^{1-\epsilon^w} + \beta \theta^w \Pi_{t+1}^{\epsilon^w-1} \Pi_t^{-\chi^w (\epsilon^w-1)} \left( \frac{W_t^*}{W_{t+1}^*} \right)^{1-\epsilon^w} F_{2,t+1} \\
\epsilon^w F_{1t} &= (\epsilon^w - 1) F_{2t}
\end{aligned}$$

Wage dispersion across unions lead to inefficiency in the labor types used by firms. This creates a wedge between hours worked  $H_t$  and effective hours worked  $H_t^d$ , which we denote by  $s_t^w$ ,

$$H_t = s_t^w H_t^d,$$

and which evolves according to,

$$s_t^w = (1 - \theta^w) \left( \frac{W_t^*}{W_t} \right)^{-\epsilon^w} + \theta \left( \frac{W_{t-1}}{W_t} \right)^{-\epsilon^w} \Pi_t^{\epsilon^w} s_{t-1}^w$$

## I.D Production of capital goods

The representative capital goods firm chooses investment  $I_t$ , the capital stock  $K_t$ , and the utilization rate  $u_t$  to maximize profits,

$$\begin{aligned}
&\max_{\{K_{t+s}, I_{t+s}, u_{t+s}\}} \sum_{s=0}^{\infty} \beta^s \lambda_{t+s} \text{Profits}_{t+s}^k \\
&\text{s.t. } \text{Profits}_t^k = R_t^k u_t K_{t-1} - I_t \\
&\quad K_t = (1 - \delta(u_t)) K_{t-1} + I_t \left[ 1 - S \left( \frac{I_t}{I_{t-1}} \right) \right]
\end{aligned}$$

where  $R_{t+s}^k$  is the rental rate of capital paid by the final goods firm,  $S \left( \frac{I_t}{I_{t-1}} \right)$  is an investment adjustment cost, and  $\delta(u)$  is the depreciation rate of capital which is increasing in utilization.

Let  $\zeta_t$  be the Lagrange multiplier on the capital accumulation equation and define Tobin's q as the relative value of capital to nondurable consumption,

$$q_t = \frac{\zeta_t}{\lambda_t^o}.$$

Then the first order conditions for the representative capital producing firms are,

$$1 = q_t \left[ 1 - S \left( \frac{I_t}{I_{t-1}} \right) - \left( \frac{I_t}{I_{t-1}} \right) S' \left( \frac{I_t}{I_{t-1}} \right) \right] + \beta \frac{\lambda_{t+1}}{\lambda_t} q_{t+1} \left( \frac{I_{t+1}}{I_t} \right)^2 S' \left( \frac{I_{t+1}}{I_t} \right)$$

$$q_t = \beta \frac{\lambda_{t+1}}{\lambda_t} R_{t+1}^k u_{t+1} + \beta (1 - \delta(u_{t+1})) \frac{\lambda_{t+1}}{\lambda_t} q_{t+1}$$

$$R_t^k = \delta'(u_t) q_t$$

## I.E Production of final goods

Final output  $Y_t$  is produced using a Cobb-Douglas production function with capital share  $\alpha$ ,

$$s_t Y_t = Z_t (u_t K_{t-1})^\alpha (H_t^d)^{1-\alpha}$$

where  $Z_t$  is aggregate TFP. The wedge  $s_t$  captures a distortion from price dispersion, which is described below.

The cost minimization for the representative final goods firm is

$$\min R_t^k u_t K_{t-1} + W_t H_t^d$$

$$\text{s.t. } Z_t (u_t K_{t-1})^\alpha (H_t^d)^{1-\alpha} = s_t Y_t$$

which yields the following first order conditions for capital and labor,

$$R_t^k = \xi_t \alpha \frac{s_t Y_t}{u_t K_{t-1}}$$

$$W_t = \xi_t (1 - \alpha) \frac{s_t Y_t}{H_t^d}$$

where  $\xi_t$  is the Lagrange multiplier on the production function. Dividing the two first order conditions yields the optimal capital-labor ratio,

$$\frac{u_t K_{t-1}}{H_t^d} = \frac{\alpha}{1 - \alpha} \frac{W_t}{R_t^k},$$

which in turn yields the marginal cost of output is,

$$MC_t = \alpha^{-\alpha} (1 - \alpha)^{-(1-\alpha)} (R_t^k)^\alpha W_t^{1-\alpha} \frac{1}{Z_t}$$

With perfect competition among final goods firms, the real final goods price is equal to marginal cost,

$$p_t^f = MC_t,$$

and final good firms make zero profits.

## I.F Prices

A continuum of retailers purchases final goods at price  $p_t^f$  and differentiates these goods with elasticity of substitution  $\epsilon$ . Retailers can only reset their price with probability  $\theta$ . The profit maximization problem for setting the reset price is

$$\max_{p_t^*} \sum_{s=0}^{\infty} \beta^s \left( \frac{\lambda_{t+s}}{\lambda_t} \right) \theta^s Y_{t+s} \left[ (p_t^*)^{1-\epsilon} \left( \frac{P_{t+s}}{P_t} \right)^{\epsilon-1} - (p_t^*)^{-\epsilon} \left( \frac{P_{t+s}}{P_t} \right)^{\epsilon} p_{t+s}^f \right]$$

The first order condition for the optimal reset price is

$$\epsilon \sum_{s=0}^{\infty} \beta^s \left( \frac{\lambda_{t+s}}{\lambda_t} \right) \theta^s Y_{t+s} \left( \frac{P_{t+s}}{P_t} \right)^{\epsilon} (p_t^*)^{-\epsilon-1} p_{t+s}^f = (\epsilon - 1) \sum_{s=0}^{\infty} \beta^s \left( \frac{\lambda_{t+s}}{\lambda_t} \right) \theta^s Y_{t+s} \left( \frac{P_{t+s}}{P_t} \right)^{\epsilon-1} (p_t^*)^{-\epsilon}$$

which we write recursively as

$$\begin{aligned} X_{1t} &= Y_t p_t^f (p_t^*)^{-\epsilon-1} + \beta \theta \left( \frac{\lambda_{t+1}}{\lambda_t} \right) \left( \frac{P_{t+1}}{P_t} \right)^{\epsilon} \left( \frac{p_t^*}{p_{t+1}^*} \right)^{-\epsilon-1} X_{1,t+1} \\ X_{2t} &= Y_t (p_t^*)^{-\epsilon} + \beta \theta \left( \frac{\lambda_{t+1}}{\lambda_t} \right) \left( \frac{P_{t+1}}{P_t} \right)^{\epsilon-1} \left( \frac{p_t^*}{p_{t+1}^*} \right)^{-\epsilon} X_{2,t+1} \\ \epsilon X_{1t} &= (\epsilon - 1) X_{2t} \end{aligned}$$

The optimal reset price determines aggregate inflation

$$1 = (1 - \theta)(p_t^*)^{1-\epsilon} + \theta \Pi_t^{-(1-\epsilon)}$$

as well as the relative price distortion

$$\begin{aligned} s_t &= \int_0^1 \left( \frac{P_t(i)}{P_t} \right)^{-\epsilon} di \\ &= (1 - \theta)(p_t^*)^{-\epsilon} + \theta \int_0^1 \left( \frac{P_{t-1}(i)}{P_t} \right)^{-\epsilon} di \\ &= (1 - \theta)(p_t^*)^{-\epsilon} + \theta \Pi_t^{\epsilon} s_{t-1} \end{aligned}$$

Due to monopoly power, the sticky-price firms make non-zero profits in equilibrium equal to

$$\text{Profits}_t^s = Y_t(1 - p_t^f)$$

## I.G Government

The central bank sets the gross nominal interest rate according to the following interest rate rule,

$$R_t = (1 - \rho_r)R_{t-1} + \rho_r \left[ R + \phi_\pi(\Pi_t - \bar{\Pi}) + \phi_y \left( \frac{Y_t}{\bar{Y}} - 1 \right) \right]$$

where  $\rho_r$  determines the degree of interest rate smoothing,  $\phi_\pi$  the response to deviations of inflation from target, and  $\phi_y$  the response to deviations of output from target.

The government issues one-period nominal bonds at gross interest  $R_t$  to cover debt repayment and any fiscal deficit.

$$B_t = \frac{R_{t-1}}{\Pi_t} B_{t-1} - T_t$$

To balance the budget over time, taxes are an increasing function of the debt level,

$$T_t = T + \phi_b(B_{t-k} - \bar{B}) - \epsilon_t.$$

We allow for a lag of  $k$  periods in the response of taxes to debt. The shock  $\epsilon_t$  represents a one-time deficit financed transfer from the government to households.

## I.H Durable Goods Production

Durable goods are produced competitively using nondurables  $N_t$  as inputs,

$$\frac{X_{it}}{p_t^d} = N_{it}$$

where  $\frac{X_{it}}{p_t^d}$  is the quantity of durable goods produced.

Real profits from the sale of durable goods are

$$\max_{N_{it}} X_{it} - N_{it} = \max_{N_{it}} p_t^d N_{it} - N_{it}$$

Profit maximization yields a flat relative supply curve,

$$p_t^d = 1.$$

## I.I Market Clearing

The goods market clears if total expenditure equals output.

$$Y_t = C_t + I_t + X_t + \eta D_t$$

The bond market clears if bonds supplied by the government equal bonds held by households,

$$B_t = A_t$$

## I.J Functional Forms

We assume the following functional forms:

$$\begin{aligned} \delta(u_t) &= \delta_0 + \delta_1(u_t - 1) + \delta_2(u_t - 1)^2 \\ S\left(\frac{I_t}{I_{t-1}}\right) &= \frac{\kappa}{2} \left(\frac{I_t}{I_{t-1}} - 1\right)^2 \end{aligned}$$

## B Additional Counterfactuals

Table B.1 shows the GE-MPCs in the model of Section V calibrated to a 2-month durable demand elasticity of -6.4 based on the estimates from Baker, Kueng, McGranahan, and Melzer (2019).

**Table B.1**  
**General Equilibrium Marginal Propensity to Consume: Model with Less Elastic Durable Demand**

PCE		Motor vehicles		Nondurable goods	
micro	GE	micro	GE	micro	GE
0.30	0.12	0.30	0.14	0.00	-0.02
0.50	0.31	0.40	0.23	0.10	0.09
0.90	1.54	0.40	0.51	0.50	1.03

## C Empirical Appendix

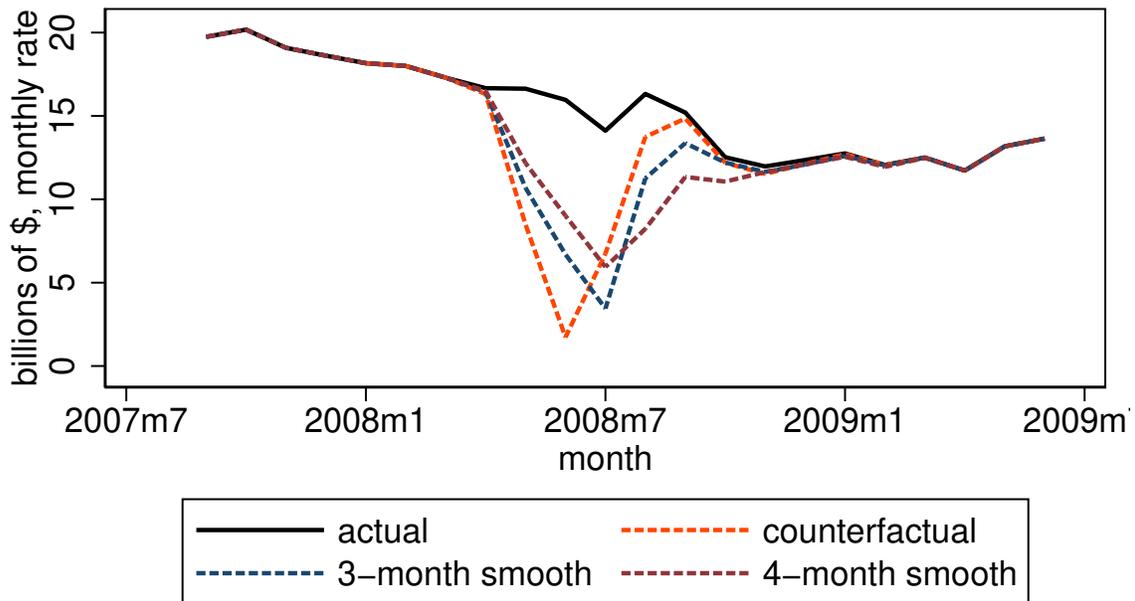
### III.A Details for Figure I

The following are details of the Sahm, Shapiro, and Slemrod (2012) calculation and our update. Sahm, Shapiro, and Slemrod (2012) use Parker, Souleles, Johnson,

and McClelland’s (2013) estimate of a marginal propensity to spend on new motor vehicles of 0.357 (from Table 7 of Parker, Souleles, Johnson, and McClelland (2013)) to calculate induced spending. Following Parker, Souleles, Johnson, and McClelland (2013), they assume that the spending is evenly distributed between the current and the next month. They use seasonal factors to seasonally adjust the induced spending. We follow the same procedure to calculate induced spending and then subtract it from actual spending to create the implied counterfactual, which does not account for partial or general equilibrium effects.

In Sahm, Shapiro, and Slemrod (2012), they assume that induced rebate spending is smoothly spent over two months, which is more concentrated than the three months we assume in our general equilibrium counterfactuals in subsection III.B. Regardless, the following graph shows that even when we use more conservative assumptions of how much spending is smoothed, the counterfactuals of motor vehicle spending are still unrealistic.

**Figure C.1**  
**Expenditures on New Motor Vehicles: Alternative Counterfactuals**



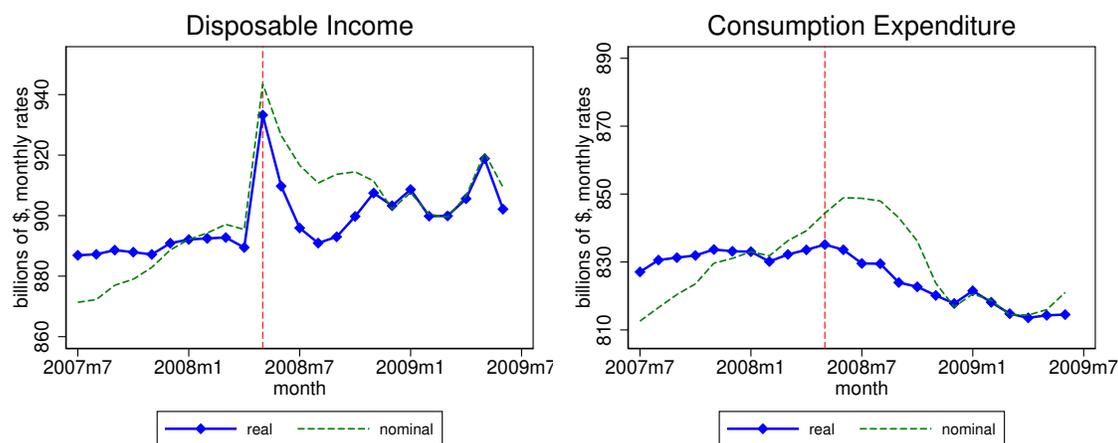
Note. The baseline counterfactual assumes that rebate-induced spending is spread over two months. The two alternatives show the counterfactual with the induced spending spread over three or four months.

### III.B Supplement to 2008 Narrative

This section supplements the 2008 narrative with details and graphs of the behavior of nominal expenditures, prices, and the federal funds rate.

Figure C.2 shows the behavior of nominal NIPA disposable personal income and consumption from mid-2007 through mid-2009, in addition to real NIPA disposable personal income and consumption that we plotted in Figure III. We normalize real income and consumption to be equal to nominal values in January 2008 for better illustration.

**Figure C.2**  
**Aggregate Disposable Income and Consumption**



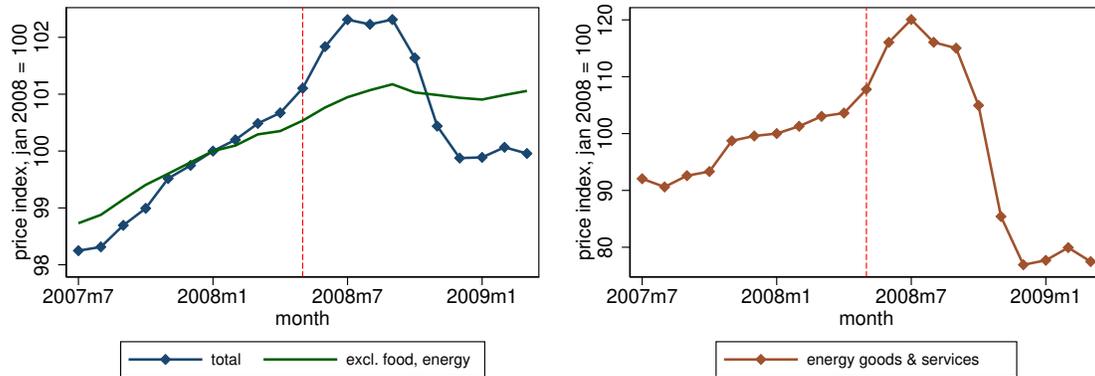
Source. BEA data. Vertical red dashed line indicates May 2008.

The effect of the 2008 tax rebate on disposable income is clearly evident in the spikes in both the nominal and real disposable income series, shown in the left panel. For both disposable income and consumption, however, the nominal and real paths look quite different from each other because of the behavior of inflation. Nominal consumption shows a prominent hump in Summer 2008, but real consumption displays only a small bump.

Figure C.3 shows the price indices for total consumption expenditures and consumption expenditures excluding food and energy. Consider first the behavior of the price deflator for total consumption. The rate of inflation for total consumption accelerated after April, resulting in July prices that were 1.6 percent above April prices. Price levels then reached a plateau and fell after the failure of Lehman Brothers in September, so that by the end of the year the level of prices was slightly lower than at the start of the year.

In contrast, the price index for consumption excluding the volatile food and energy components showed a more modest rate of inflation, averaging 3.4 percent annualized

**Figure C.3**  
**Price Indices for Consumption**



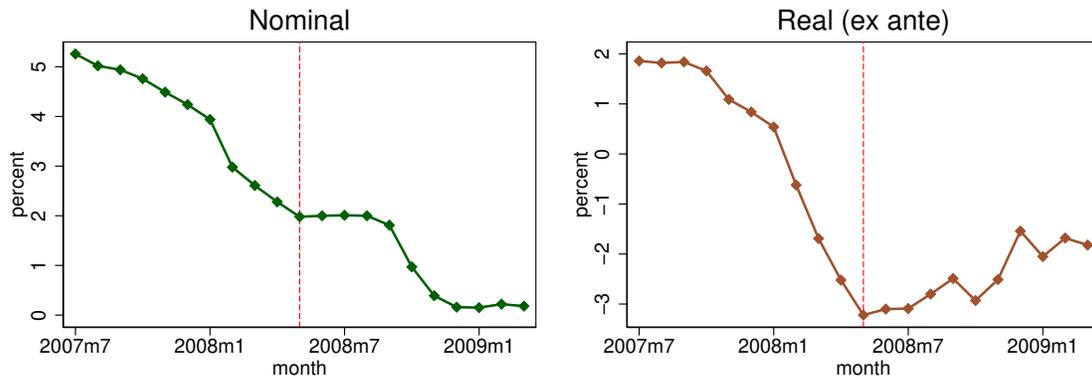
Source. BEA data.

for January through the peak in September 2008. This price level then declined slightly after the collapse of Lehman Brothers.

A key source of volatility of consumer prices in 2008 was the behavior of crude oil prices (not shown). The price for West Texas Intermediate rose from \$98 per barrel in January 2008 to a peak of \$140 per barrel in July 2008. By the end of 2008, it had fallen to only \$33 per barrel.

Turning to interest rates, Figure C.4 shows the behavior of the nominal and ex ante real federal funds rate. The ex ante real federal funds rate is the difference between the nominal federal funds rate and the current month median expected annual inflation rate from the University of Michigan Survey of Consumers. The nominal series shows cuts every month from mid-2007 to May 2008, a leveling off from May through August, and then cuts until the zero lower bound was reached. The combination of the cuts and the higher expected rates of inflation result in negative real interest rates starting in February 2008.

**Figure C.4**  
**Federal Funds Rate**



Source. FRED, based on Federal Reserve Board of Governors. The ex ante real interest rate is constructed using the current month median expected annual inflation rate from the University of Michigan Survey of Consumers.

### III.C Alternative Measures of Consumption Expenditures

Because the monthly NIPA consumption data are based on combining and smoothing various data sources, we provide supplemental evidence that the patterns we showed for consumption expenditures in Figure III are not due to smoothing procedures.

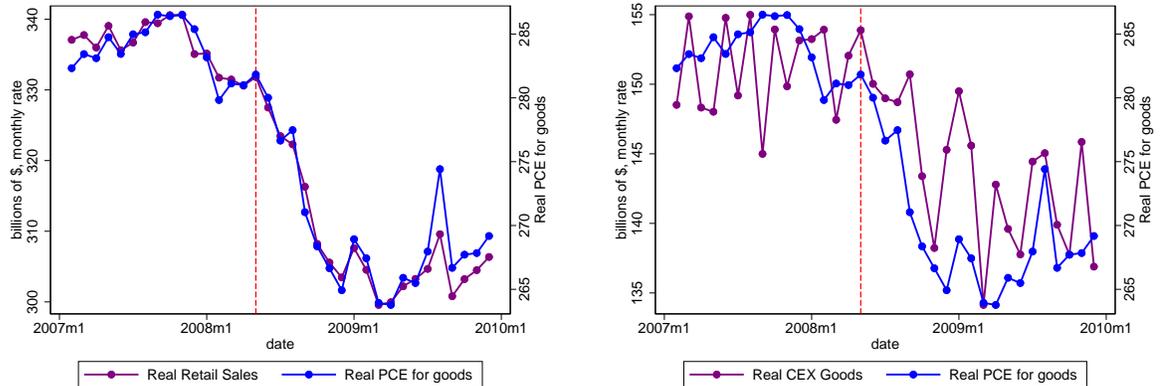
We compare the NIPA measures of personal consumption expenditures (PCE) on goods to two other series: the Census series on retail sales of goods and our own constructed series based on the CEX data that is the basis for the micro estimates.<sup>1</sup> As described by Wilcox (1992), government statisticians use retail sales as an input to monthly consumption, but then make a number of adjustments. To make sure those adjustments are not smoothing out jumps in consumption due to the rebate, we examine the key underlying series as well as our constructed alternative from the CEX. For all series, we use the PCE goods deflator to create real spending series.

Figure C.5 shows the spending comparisons from 2007 through 2009. Consider first the left side graph, which compares PCE on goods to retail sales. The movements in the two series match up very well over the two years. Both show a slight blip up in May 2008, with the retail series showing a more muted blip. Thus, it is unlikely that BEA smoothing of retail sales would account for the consumption pattern.

The right-hand side graph compares PCE on goods to our aggregates of household spending on goods using CEX micro data. The CEX aggregate is much noisier than either the PCE data or the retail sales data. The CEX series falls from February to March,

1. To construct this series we match categorical spending in the CEX to NIPA spending following the concordance prepared by the BLS staff (Bureau of Labor Statistics, 2019).

**Figure C.5**  
**Comparison of PCE to Retail and CEX**



Source. PCE (Personal Consumption Expenditures) from BEA; Retail sales from Census; Authors' aggregation from CEX. Vertical red dashed line indicates May 2008.

recovers in April, and then declines in May and June. These movements look similar to those in other months, suggesting more noise than information.

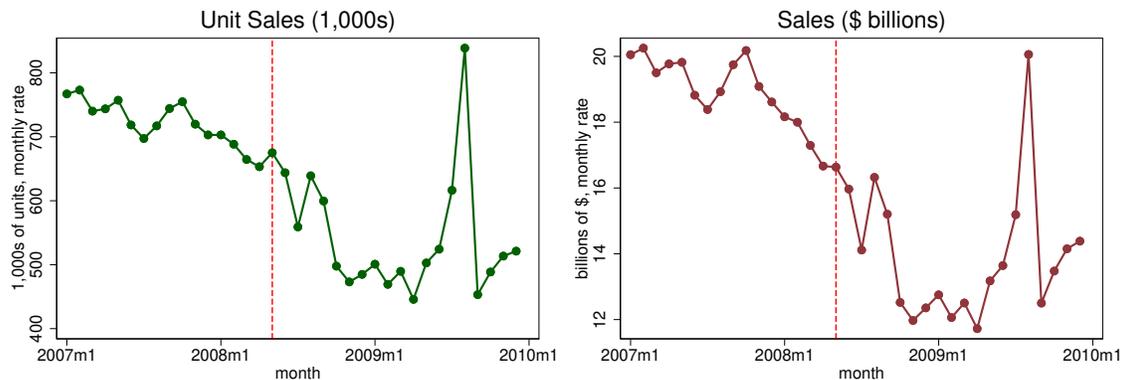
In Figure C.6 we plot new motor vehicle sales in units (left panel) and in dollars (right panel). Due to mandatory registration new motor vehicle sales are essentially perfectly measured. They are also one of the categories in which Parker, Souleles, Johnson, and McClelland (2013) estimate large MPCs. Thus, these data should be informative whether there is an expenditure spike following the 2008 rebate. The unit sales in the left panel show a small blip in May 2008 (red vertical line), but there is little change in dollar sales in the right panel. Neither data series shows a large spike in motor vehicle purchases around the time the rebate is received by households.

We conclude that the PCE data is not smoothing out a large jump in consumption when the rebates are distributed.

### III.D Forecast Appendix

Our forecasting model is a simple monthly-frequency time series model with the following endogenous variables: log real consumption, log real disposable income, log consumption deflator, and the Gilchrist and Zakrajšek (2012) excess bond premium. We also include a dummy variable for recession, log real oil prices, and a dummy variable for the Lehman Brothers bankruptcy in September 2008. We use six lags of each variable. We include current values of the recession dummy, oil prices, and the excess bond premium in the equations for the endogenous variables. We estimate the model on data from 1984m1 - 2019m12 and forecast dynamically starting in January 2008. We start the estimation period in 1984 because the effects of oil prices on consumption expenditures changed significantly post-1984 (e.g. Edelstein and Kilian (2009)).

**Figure C.6**  
**New Motor Vehicle Sales**



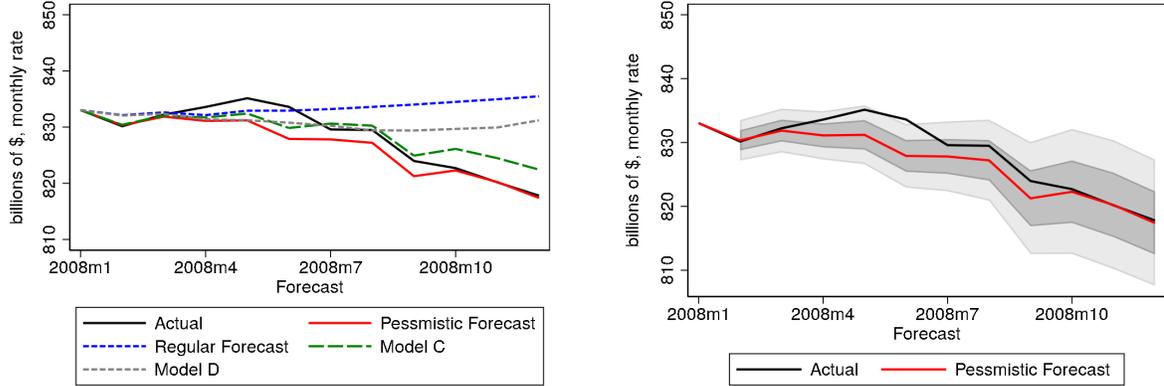
Source: Unit sales and total sales from BEA.

We produce four forecasts by varying our assumptions on whether oil prices followed their actual path and whether the recession and Lehman Brothers dummies. The most pessimistic forecasts are those in which oil prices are assumed to follow their actual path and in which the recession and Lehman Brothers bankruptcy dummy variables are included in the forecasting equation. The regular forecast excludes the recession and Lehman dummies and models oil prices as endogenous. Model C includes the recession dummies and models oil prices as endogenous, whereas model D excludes the dummies and models oil prices as exogenous. Figure C.7 show that models C and D are spanned by the pessimistic and regular forecast.

### III.E Vehicle Demand Elasticity from Mian and Sufi (2012)

The estimates in Mian and Sufi (2012) from the 2009 Cash-for-Clunkers program imply a two-month demand elasticity ranging from -26 to -30 compared to -26 in our model. Mian and Sufi (2012) argue that cross-city variation in Cash for Clunkers explains between 340k and 398k of additional autos sold in July and August. New vehicle sales in April and May were on average 833,000. Used vehicle sales were 36.5m and 35.5m in 2008 and 2009, implying an average monthly sales volume of 2m. Total baseline vehicle sales over two months are then 5.666m. The increase in vehicle sales estimated by Mian and Sufi then corresponds to  $340k/5.666m = 6\%$  to  $398k/5.666m = 7\%$  rise. Total expenditure on Cash for Clunkers was \$3bn, and the vehicle stock in 2008 was worth \$1279.4bn at replacement cost. This translates into a  $3/1279.4 = 0.23$  percent reduction in the replacement price. Therefore, the elasticity implied by these estimates ranges from  $-6/0.23 = -26$  to  $-7/0.23 = -30$ .

**Figure C.7**  
**Our Real Consumption Forecasts**



Alternate Forecasts

Pessimistic Forecast with Confidence Interval

Forecasts are based on information through January 2008, with exception of models in which oil prices are assumed to follow their actual path and Lehman Brothers dummies are included. Real oil prices exogenously follow their actual path in the pessimistic forecast and model D; Recession and Lehman Brothers bankruptcy dummy variables are included in the pessimistic forecast and model C. The right panel shows the 1-standard deviation (dark gray) and 2-standard deviation confidence intervals for the pessimistic forecast.

### III.F Anticipation Effects

One potential issue in identifying the micro MPCs is pre-anticipation on the part of rebate recipients. For example, households could anticipate receiving a rebate and lower/raise their spending in the months prior to receipt. One common method to test for pre-trends is to include a lead of the treatment variable in an OLS two-way fixed effect specification, however, this type of estimation results in similar negative weighting and forbidden comparisons as estimating the treatment itself using OLS (Borusyak, Jaravel, and Spiess, 2024).

Following Borusyak, Jaravel, and Spiess (2024) we test for anticipation effects by restricting our sample to households that never receive a rebate or who have not yet received a rebate and we modify equation (17) to include an indicator,  $I(ESP_{i,t+2})$ , equal to one if the household receives a rebate in the following period:

$$C_{i,t+1} - C_{i,t} = \sum_s \beta_{0s} month_{s,i} + \beta'_1 \mathbf{X}_{i,t} + \gamma I(ESP_{i,t+2}) + \tilde{u}_{i,t+1}, \quad \forall (i, t + 1) \in \{\text{Untreated}\}$$

Finally, we test whether  $\hat{\gamma} = 0$  using an F-Test. Table C.1 shows the estimated coefficients and the corresponding F-Tests for each of our main regression specifications along

**Table C.1**  
**Test for Pre-trends following BJS**

	Full Sample			Rebate Only Sample		
	PCE (1)	Vehicles (2)	Other (3)	PCE (4)	Vehicles (5)	Other (6)
Pre-trend	−186.0 (218.0)	17.5 (146.4)	−203.5 (169.0)	−392.9 (460.5)	−111.8 (321.3)	−281.1 (337.2)
F-Stat	0.73	0.01	1.45	0.73	0.12	0.69
P-Value	(0.39)	(0.90)	(0.23)	(0.39)	(0.73)	(0.40)
Observations	9573	9573	9573	2687	2687	2687

Notes: Pre-trend coefficients computed following section 4.4 of Boryusak, Jaravel, and Spiess (2022) (BJS). Standard errors, in parentheses, are clustered at the household level. Significance is indicated by: \*  $p < 0.1$ , \*\*  $p < 0.05$ , \*\*\*  $p < 0.01$ . The null hypothesis for the F-test is that the coefficient on the pre-treatment indicator is equal to zero. All regressions include interview (time) fixed effects, as well as household level controls for age, change in number of adults, and change in number of children. Extra controls refer to additional controls for household income decile and lagged total spending.

with the regressions including additional controls. We cannot reject the assumption of no anticipation effects.

### III.G Model Regressions

In this section we simulate data from the model in section V and repeat our empirical approach in section IV. Specifically, we simulate data that has the same overlapping structure and the same distribution of rebates as the CEX. Table C.2 shows that the baseline specification in Parker, Souleles, Johnson, and McClelland (2013) (17) does not recover the true MPCs because it does not account for the comparison with previously treated units. By contrast, our preferred specification in column 4 of Table III correctly recovers the underlying MPC. Table C.3 shows that the bias of TWFE in equation (17) gets worse in the rebate only sample because that sample makes relatively more comparisons with previously treated units.

**Table C.2**  
**Rebate Coefficient MPC Estimates in Model: Full Sample**

Dependent Variable: Change in Total Expenditures						
Specification:	Table III, Column 1			Table III, Column 4		
True MPC:	0.3	0.5	0.9	0.3	0.5	0.9
Rebate Coeff.	334.1	556.9	1002.0	286.0	476.6	857.4
Implied MPC	0.35	0.59	1.05	0.30	0.50	0.90

Notes: The dependent variable is the change in total expenditures from the previous interview. The rebate size in the model is \$950 for all households conditional on receiving a rebate. The rebate only sample includes only households that receive a rebate at some point during our sample period.

**Table C.3**  
**Rebate Coefficient MPC Estimates in Model: Rebate Only Sample**

Dependent Variable: Change in Total Expenditures						
Specification:	Table III, Column 1			Table III, Column 4		
True MPC:	0.3	0.5	0.9	0.3	0.5	0.9
Rebate Coeff.	426.3	710.8	1278.3	286.1	476.6	857.3
Implied MPC	0.45	0.75	1.35	0.30	0.50	0.90

Notes: The dependent variable is the change in total expenditures from the previous interview. The rebate size in the model is \$950 for all households conditional on receiving a rebate. The rebate only sample includes only households that receive a rebate at some point during our sample period.

### III.H Comparison to Borusyak, Jaravel, and Spiess (2024) Method

A leading alternative approach to accounting for the bias from omitting lagged treatment variables and treatment effect heterogeneity is the imputation method of Borusyak, Jaravel, and Spiess (2024). Their method imposes a linear fixed effect structure (parallel trends) only on the untreated potential outcomes. In particular, their method begins by estimating potential outcomes for untreated units in order to make proper control group comparisons. They estimate

$$(36) \quad Y_{i,t}(0) \equiv C_{i,t}(0) - C_{i,t-1} = \sum_s \beta_{0,s} month_s + \beta_1' \mathbf{X}_{i,t} + \beta_4 C_{i,t-1} + u_{i,t}$$

where  $C_{i,t}(0)$  is consumption of household  $i$  if it were not treated at  $t$  and  $Y_{i,t}(0)$  is the associated change in consumption (so the treatment indicator is 0). Since estimation is only for untreated units, there are no *ESP* terms in this specification.

If there are no anticipation effects,  $Y_{i,t} = Y_{i,t}(0)$  for the untreated and not-yet treated units. The treatment effect for a household  $i$  who receives the rebate at  $t$  is then  $Y_{i,t} -$

$Y_{i,t}(0)$ . The aggregate contemporaneous and lagged treatment effects are:

$$(37) \quad \tau_0 = \sum_i \omega_i [Y_{i,E(i)} - Y_{i,E(i)}(0)], \quad E(i) : \{\text{Treatment date of household } i\}$$

$$(38) \quad \tau_1 = \sum_i \omega_i [Y_{i,E(i)+1} - Y_{i,E(i)+1}(0)]$$

for household weights  $\omega_i$ .

A significant advantage of the BJS method over our OLS specification is that it does not impose a particular dynamic structure of the treatment effects, since it only imposes structure on untreated outcomes in (36). For example, if the rebate affected the change in consumption beyond one lag as we imposed in (32), then the BJS estimator remains consistent whereas OLS is biased. Furthermore, the BJS estimator imposes no restriction on the form of heterogeneity in the treatment effects. Thus, the BJS estimator overcomes the first two biases we identify—the omitted variable bias and the forbidden comparison bias—by construction. Since the BJS estimator imposes weaker assumptions than OLS it is more likely to be a consistent estimator of the average treatment effect in our setting. However, it is less efficient if the model we imposed for our OLS estimator in (32) is true.<sup>2</sup>

Appendix Table C.4, columns (2) and (4), show that the treatment effects from the BJS procedure are very similar to our final estimates in column (4) of Table III.<sup>3</sup> The MPC in the full sample is slightly lower at 0.2 and the MPC in the rebate-only sample is slightly higher at 0.37. The average of the two is essentially the same as the average of our OLS estimates. Thus, allowing for more general dynamic treatment effects and heterogeneity than our final estimation equation (32) has only small effects on the implied MPCs.

### III.I Estimated MPCs for Motor Vehicles and Parts and Other PCE

Tables C.5 through Table C.8 show estimated MPCs for motor vehicles and parts, and for the residual PCE expenditure (“other PCE”) that are referenced in the paper.

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2. Even when the model is true, the OLS and BJS estimators will generally produce different estimates for the average treatment effect because BJS weighs treated observations using sample weights, whereas OLS will base weights also on the variance of the treatment indicator over time.

3. We use Borusyak, Jaravel, and Spiess (2024)’s `did_imputation` STATA command to construct point estimates and standard errors.

**Table C.4**  
**Contemporaneous Household PCE Response to Rebate: BJS Method**

	Full Sample		Rebate Only Sample	
	(1)	(2)	(3)	(4)
Rebate Indicator	294.2 (221.5)	191.3 (193.3)	917.0 (667.8)	351.0 (557.4)
Lag Total Expenditure		-0.25*** (0.039)		-0.30*** (0.050)
Lag Motor Vehicle		-0.75*** (0.046)		-0.71*** (0.057)
Implied MPC	0.31	0.20	0.96	0.37
Income Decile FE	No	Yes	No	Yes
Observations	12,425	12,425	5,511	5,511

Notes: The dependent variable is the change in PCE. Regressions include interview (time) fixed effects, and household level controls for age, change in number of adults, and change in number of children. Standard errors, in parentheses, are clustered at the household level. Significance is indicated by: \*  $p < 0.1$ , \*\*  $p < 0.05$ , \*\*\*  $p < 0.01$ .

**Table C.5**  
**First Stage: Rebate Amount Conditional on Rebate Receipt**

<b>Panel A: Full Sample</b>				
	Homogeneous Treatment		Heterogeneous Treatment	
	(1)	(2)	(3)	(4)
Rebate Indicator	948.60*** (10.37)	951.10*** (10.29)	950.54*** (10.19)	945.95*** (10.07)
Lag Rebate Indicator		11.97*** (3.17)	1.19** (0.60)	-2.00 (1.29)
Lag Total Expenditure				0.00*** (0.00)
Lag Motor Vehicle				-0.00 (0.00)
Income Decile FE	No	No	No	Yes
Observations	16,962	16,962	16,962	16,962
<b>Panel B: Rebate Recipients Only</b>				
	Homogeneous Treatment		Heterogeneous Treatment	
	(1)	(2)	(3)	(4)
Rebate Indicator	931.69*** (13.11)	945.06*** (12.73)	939.42*** (12.53)	946.84*** (13.25)
Lag Rebate Indicator		24.14*** (7.73)	-1.14 (2.08)	10.44** (4.63)
Lag Total Expenditure				0.00*** (0.00)
Lag Motor Vehicle				-0.00*** (0.00)
Income Decile FE	No	No	No	Yes
Observations	10,076	10,076	10,076	10,076

Notes: The dependent variable is the dollar value of Economic Stimulus Payments (ESP) received by the household. Standard errors, in parentheses, are clustered at the household level. Regressions include interview (time) fixed effects, and household level controls for age, change in number of adults, and change in number of children. The rebate coefficients in columns (3) and (4) are the weighted average of the interaction between rebate cohort and the (lagged) rebate indicator with weights computed following Sun and Abraham (2021). Significance is indicated by: \*  $p < 0.1$ , \*\*  $p < 0.05$ , \*\*\*  $p < 0.01$ .

**Table C.6**  
**Household Total Spending (PSMJ) Response to Rebate**

<b>Panel A: Full Sample</b>					
	Homogeneous Treatment		Heterogeneous Treatment		
	(1)	(2)	(3)	(4)	(5)
Rebate Indicator	483.19** (209.87)	417.87** (202.02)	325.96 (203.45)	251.71 (178.69)	142.94 (176.91)
Lag 1 Rebate Indicator		-377.83* (214.64)	-460.14** (220.78)	-307.98* (175.37)	-313.88* (180.96)
Lag Total Spending (PSMJ)				-0.50*** (0.04)	-0.50*** (0.04)
Lag Motor Vehicle (PSMJ)				-0.50*** (0.05)	-0.51*** (0.05)
Lag Non-durable (PSMJ)				0.36*** (0.06)	0.37*** (0.06)
Implied 3-month MPC	0.52	0.45	0.35	0.27	0.15
Income Decile FE	No	No	No	Yes	Yes
Exclude 2+ Rebate	No	No	No	No	Yes
Observations	17,229	17,229	17,229	17,229	16,962
<b>Panel B: Rebate Recipients Only</b>					
	Homogeneous Treatment		Heterogeneous Treatment		
	(1)	(2)	(3)	(4)	(5)
Rebate Indicator	779.23** (310.22)	551.34* (315.93)	-276.02 (447.00)	266.66 (312.34)	122.10 (567.69)
Lag 1 Rebate Indicator		-462.55 (330.43)	-1316.29*** (508.51)	-541.04* (296.34)	-450.63 (456.07)
Lag Total Spending (PSMJ)				-0.56*** (0.04)	-0.56*** (0.04)
Lag Motor Vehicle (PSMJ)				-0.44*** (0.04)	-0.45*** (0.04)
Lag Non-durable (PSMJ)				0.41*** (0.07)	0.42*** (0.07)
Implied 3-month MPC	0.86	0.61	-0.31	0.30	0.13
Income Decile FE	No	No	No	Yes	Yes
Exclude 2+ Rebate	No	No	No	No	Yes
Observations	10,343	10,343	10,343	10,343	10,076

Notes: The dependent variable is the change in Total Spending (PSMJ). Regressions include interview (time) fixed effects, and household level controls for age, change in number of adults, and change in number of children. Standard errors for the 6-month MPC are estimated via Delta-method. The rebate coefficients in columns (3), (4), and (5) are the weighted average of the interaction between rebate cohort and the (lagged) rebate indicator with weights computed following Sun and Abraham (2021). Standard errors, in parentheses, are clustered at the household level: \*  $p < 0.1$ , \*\*  $p < 0.05$ , \*\*\*  $p < 0.01$ .

**Table C.7**  
**Household Motor Vehicles (PSMJ) Response to Rebate**

<b>Panel A: Full Sample</b>					
	Homogeneous Treatment		Heterogeneous Treatment		
	(1)	(2)	(3)	(4)	(5)
Rebate Indicator	371.80** (154.33)	344.36** (147.19)	307.02** (147.06)	397.73*** (121.19)	284.85*** (107.41)
Lag 1 Rebate Indicator		-158.69 (154.50)	-225.49 (155.18)	87.22 (103.25)	75.20 (104.96)
Lag Total Spending (PSMJ)				0.00 (0.01)	0.00 (0.01)
Lag Motor Vehicle (PSMJ)				-1.01*** (0.01)	-1.01*** (0.01)
Lag Non-durable (PSMJ)				0.04** (0.02)	0.04** (0.02)
Implied 3-month MPC	0.40	0.37	0.33	0.43	0.30
Income Decile FE	No	No	No	Yes	Yes
Exclude 2+ Rebate	No	No	No	No	Yes
Observations	17,229	17,229	17,229	17,229	16,962
<b>Panel B: Rebate Recipients Only</b>					
	Homogeneous Treatment		Heterogeneous Treatment		
	(1)	(2)	(3)	(4)	(5)
Rebate Indicator	389.29* (231.18)	273.37 (229.01)	-428.58 (371.84)	198.38 (178.29)	326.78** (157.26)
Lag 1 Rebate Indicator		-235.30 (255.14)	-952.16** (455.38)	-216.37 (164.57)	103.29 (160.30)
Lag Total Spending (PSMJ)				-0.01 (0.01)	-0.01 (0.01)
Lag Motor Vehicle (PSMJ)				-0.99*** (0.02)	-1.01*** (0.01)
Lag Non-durable (PSMJ)				0.06** (0.02)	0.06** (0.02)
Implied 3-month MPC	0.43	0.30	-0.48	0.22	0.34
Income Decile FE	No	No	No	Yes	Yes
Exclude 2+ Rebate	No	No	No	No	Yes
Observations	10,343	10,343	10,343	10,343	10,076

Notes: The dependent variable is the change in Vehicle Spending (PSMJ). Regressions include interview (time) fixed effects, and household level controls for age, change in number of adults, and change in number of children. Standard errors for the 6-month MPC are estimated via Delta-method. The rebate coefficients in columns (3), (4), and (5) are the weighted average of the interaction between rebate cohort and the (lagged) rebate indicator with weights computed following Sun and Abraham (2021). Standard errors, in parentheses, are clustered at the household level: \*  $p < 0.1$ , \*\*  $p < 0.05$ , \*\*\*  $p < 0.01$ .

**Table C.8**  
**Monthly: Household Motor Vehicle and Parts Spending Response to Rebate**

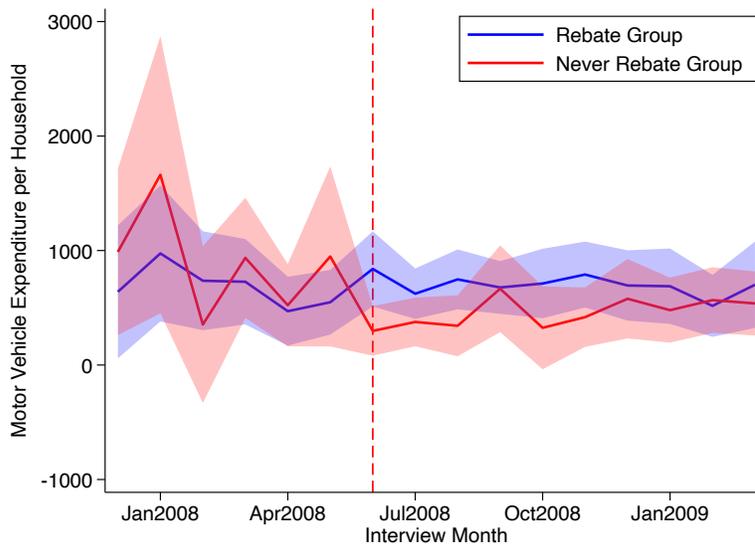
	Full Sample	Rebate Only Sample
	(1)	(2)
Lead 1 Rebate Indicator	132.87** (59.46)	110.68* (66.27)
Rebate Indicator	75.40 (66.94)	54.56 (78.28)
Lag 1 Rebate Indicator	39.98 (65.82)	30.06 (77.13)
Lag 2 Rebate Indicator	75.68 (58.43)	72.06 (64.68)
Lag 3 Rebate Indicator	33.53 (60.93)	26.38 (69.46)
Lag Total Expenditure	0.01 (0.01)	0.03* (0.01)
Lag Motor Vehicle	-1.16*** (0.01)	-1.15*** (0.02)
Income Decile FE	Yes	Yes
Observations	44,996	26,418

Notes: The dependent variable is the change in motor vehicles and parts expenditure. Regressions include month fixed effects, and household level controls for age, change in number of adults, and change in number of children. The rebate coefficients are the weighted average of the interaction between rebate month cohort and the rebate indicators with weights computed following Sun and Abraham (2021). Standard errors, in parentheses, are clustered at the household level: \*  $p < 0.1$ , \*\*  $p < 0.05$ , \*\*\*  $p < 0.01$ .

### III.J Motor Vehicle Spending by Rebate Status

An implication of the crowding out mechanism is that the control group in the household regressions in Section IV is also affected by the rebate. Specifically, the group of households that does not receive a rebate will reduce its expenditures on motor vehicles because of the temporarily higher motor vehicle price.<sup>4</sup> Figure C.8 plots suggestive evidence for this mechanism in the CEX data. Motor vehicle expenditure rises in the treated group when the first rebates are reported in June 2008, but it simultaneously falls in the control group. The differences in motor vehicle expenditure in June is statistically significant at the 1% level. However, we cannot rule out that another macro shock could have reduced motor vehicle expenditure exactly in June 2008.

**Figure C.8**  
**Motor Vehicle Spending per Household by Rebate Status**



Source: CEX and author's calculation. The rebate group is the set of in-sample households that ever report receiving a rebate. The never rebate group is the set of in-sample households that never report receiving a rebate. The red dashed line is June 2008, the first interview month in which expenditure and rebate receipt for May 2008 are recorded.

4. There is also a symmetric crowding out effect for the group of households that do receive rebates, so that the difference in spending—the micro MPC—is unaffected and only the aggregate GE-MPC falls.

## D Monte Carlo: Lag Expenditure Correlated with Rebate Report

In this section we show Monte Carlo simulations where households selectively report receiving a rebate based on lagged expenditure. Selective reporting of the rebate leads to an upward biased rebate coefficient because the expenditure process is mean reverting, so selective reporting on lagged expenditure creates a correlation between rebate reporting and the change in household expenditure. The upward bias declines when we condition on lagged expenditure. Results from the Monte Carlo simulations are able to quantitatively explain both the correlation between lagged expenditure and the rebate coefficient that we find in Table 4, as well as the drop in the rebate coefficient after including lagged expenditure as a control that we find in Table 3.

### IV.A Data Generating Process

The data generating process for each of our Monte Carlo runs is as follows. We simulate 10000 runs.

We assume that the true household expenditure process without rebates follows an AR(1):

$$(39) \quad \tilde{C}_{it} = \rho \tilde{C}_{i,t-1} + \epsilon_{it}, \quad \epsilon_{it} \sim N(0, \sigma_c^2)$$

We also assume that the rebate induces a spike in expenditure of size  $\beta$  upon receipt,

$$(40) \quad C_{it} = \tilde{C}_{it} + \beta I(ESP_{i,t})$$

where  $I(ESP_{i,t})$  is an indicator function equal to one if household  $i$  receives a rebate at time  $t$ .

We calibrate the persistence  $\rho$  and the standard deviation of the expenditure shock  $\sigma_c$  to our estimates from the CE data in 2008. This yields  $\rho = 0.68$  and  $\sigma_c = \$7700$ . The mean is irrelevant in our analysis so we normalize it to zero. Based on our preferred estimates we set the effect of the rebate on consumption to  $\beta = \$300$ . In each of our Monte Carlo runs, we simulate consumption expenditures for 8000 households each with four observations. Households may receive the rebate in periods two, three, and four.

We assume that both the assignment of the rebate  $I(ESP_{i,t})$  and the timing across the three periods is random. Specifically, a random 85% of households will receive the rebate over the three periods (Parker, Souleles, Johnson, and McClelland, 2011). Then, among the recipients, a random one third will receive the rebate in each of the periods two, three, and four. A household will receive at most one rebate in our simulation.

## IV.B Misreporting Rebate Receipt

True rebate receipt  $I(ESP_{i,t})$  is not observed in the CEX. We only observe reported rebate receipt, which we denote  $I(ESP^*_{i,t})$ . We refer to the difference between the two as “misreporting.”

The following pieces of evidence suggest that misreporting is present in the CEX:

1. Only 70% of households in the CEX that are sampled continuously from June through November report receiving a rebate, compared to the 85% of tax units that are eligible for a rebate.<sup>5</sup>
2. Figure 1 in Parker, Souleles, Johnson, and McClelland (2011) shows that rebate reporting in the CEX lags Treasury payout data.
3. Table D.1 shows that households disproportionately report receiving a rebate in the month prior to the interview. See Section E for how this table is constructed.

Misreporting can take different forms. As a baseline case we assume that only a random fraction  $1 - p_2$  of rebates are being reported. Thus  $p_2$  is fraction of false negative reports (Type-II error). We pick the false negative rate  $p_2$  to match the ratio of the fraction of households in our sample that report receiving a rebate (0.7) over the fraction of tax units eligible for the rebate (0.85).

We show results for other specifications of misreporting in Section IVF.

## IV.C Why Selection?

We argue that the misreporting error  $I(ESP^*_{i,t}) - I(ESP_{i,t})$  is not random (classical), but that there is selection in misreporting. We base this claim on the following observations:

1. Table 3 shows that including lagged expenditure reduces the rebate coefficient, which should not occur if rebate reporting is random.
2. Table 4 shows that low lagged expenditure predicts future rebate reporting.
3. Table D.2 shows that households that report receiving the rebate in the prior month report a much greater increase in expenditure than households that report receiving the rebate three months ago. See Section VB for how this table is constructed.

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5. Households and tax units can coincide, but there are some types of tax filers such as married filing separately that would qualify as two tax units, but one household.

**Table D.1**  
**Distribution of CEX Interview Schedule**

Panel A: EFT and Check Recipients				
Interview Schedule	Overall CEX	May Cohort	June Cohort	July Cohort
Jan-Apr-Jul-Oct	33%	32%	35%	26%
Feb-May-Aug-Nov	34%	29%	37%	39%
Mar-Jun-Sep-Dec	34%	39%	28%	34%

Panel B: Check Recipients Only			
Interview Schedule	May Cohort	June Cohort	July Cohort
Jan-Apr-Jul-Oct	30%	36%	28%
Feb-May-Aug-Nov	34%	35%	40%
Mar-Jun-Sep-Dec	36%	28%	32%

Notes: Data in column 1 come from the entire CEX Sample 2007-2009. Data in columns 2-4 come from our subsample. See Section VA for details.

**Table D.2**  
**PCE Response to Rebate:**

	Full Sample			Rebate Only		
	(1)	(2)	(3)	(4)	(5)	(6)
Right hand side variables:						
Interview 1 month post rebate	735.0*** (273.0)	722.5*** (271.2)	1022.0* (565.8)	758.5** (376.2)	721.9* (377.2)	786.1 (578.8)
Interview 2 months post rebate	413.2 (301.0)	387.8 (303.8)	1097.3 (894.7)	587.2 (427.4)	536.9 (448.2)	1025.8 (902.9)
Interview 3 months post rebate	112.0 (299.3)	68.0 (309.8)	-437.8 (719.1)	174.9 (409.5)	107.5 (451.9)	-594.5 (697.6)
Interview 4 months post rebate		-406.9 (316.3)	-285.0 (255.3)		-613.3 (471.5)	-301.8 (360.8)
Interview 5 months post rebate		-322.5 (307.1)	-342.9 (255.2)		-738.0* (432.0)	-560.3 (346.4)
Interview 6 months post rebate		42.7 (329.1)	170.1 (244.6)		-391.3 (404.0)	-69.7 (292.2)
One month = Two month p-value	0.35	0.34	0.94	0.63	0.61	0.80
One month = Three month p-value	0.08	0.08	0.08	0.12	0.12	0.10
Time Effects	Yes	Yes	Yes	Yes	Yes	Yes
Household Controls	Yes	Yes	Yes	Yes	Yes	Yes
Income Decile FE	No	No	Yes	No	No	Yes
Lag Expenditure Interacted Rebate Month	No	No	Yes	No	No	Yes
Observations	16,962	16,962	16,962	10,076	10,076	10,076

Notes: The dependent variable is the change in Total Spending (PSMJ). Regressions include interview (time) fixed effects, and household level controls for age, change in number of adults, and change in number of children. Standard errors, in parentheses, are clustered at the household level: \*  $p < 0.1$ , \*\*  $p < 0.05$ , \*\*\*  $p < 0.01$ . See Section V.B for details.

We assume selection is a function of lagged expenditure  $C_{i,t-1}$ . We do so because Table 4 shows that expenditure is unusually low in the interview before the rebate was reported received.

We model selection into reporting a rebate based on a latent variable

$$(41) \quad y_{it}^* = \Phi^{-1}(1 - p_2)I(ESP_{i,t}) - \gamma C_{i,t-1} + u_t, \quad u_t \sim N(0, 1)$$

where  $\Phi^{-1}(\cdot)$  is the inverse of a standard normal,  $\gamma C_{i,t-1}$  captures selection based on lagged expenditure, and  $u_t$  is drawn from a standard normal.

We define an indicator that the household reports receiving a rebate as:

$$(42) \quad I(ESP^*_{i,t}) = 1(y_{it}^* > G(p_2)) \times I(ESP_{i,t})$$

where  $G(\cdot)$  is the inverse CDF of the distribution of the latent variable amongst rebate recipients.<sup>6</sup>

The parameter  $\gamma$  controls the strength of selection. When  $\gamma = 0$ , misreporting is random so the measurement error in  $I(ESP^*_{i,t})$  is classical. When  $\gamma > 0$ , households with lower lagged expenditure are more likely to report a rebate in the following interview. Since expenditure is a mean reverting process, low lagged expenditure today forecasts a larger change in expenditure tomorrow. Thus selection on lagged expenditure induces a correlation between a household's change in expenditure and their probability to report receiving a rebate, which biases the estimated rebate coefficient as we show next.

#### IV.D Disciplining the Strength of Selection $\gamma$

We discipline the strength of selection  $\gamma$  by targeting the lead rebate coefficient  $\zeta$  in Table 4 of the paper. That Table is based on the regression,

$$(43) \quad C_{it} = \xi I(ESP^*)_{it} + \zeta I(ESP^*)_{i,t+1} + v_{it}$$

where the lead coefficient  $\zeta$  captures the degree to which low current expenditure predicts future rebate receipt. We focus on the coefficient for the rebate only sample, because in that sample the true rebate assignment should be random,  $E[I(ESP)_{it} v_{it}] = E[I(ESP)_{i,t+1} v_{it}] = 0$ . However, in the data we estimate  $\zeta_x = -575$ , which suggests that rebates are disproportionately reported for households with unusually low lagged expenditure.

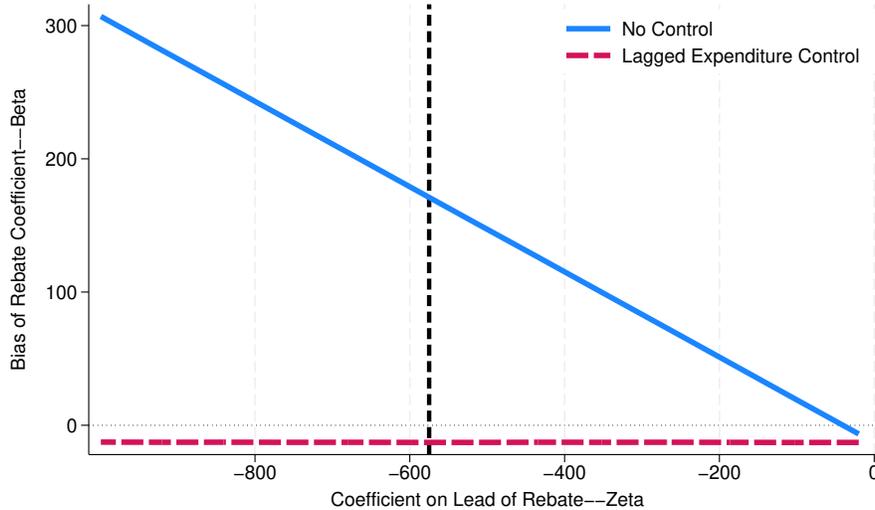
#### IV.E Quantitative Assessment of the Selection Bias

Figure D.1 maps the MPC bias into the lead coefficient  $\zeta$ . The mapping is monotonic: the larger the lead coefficient, the larger is the degree of selection, and the larger is the

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6.  $G(p_2) = 0$  when  $\gamma = 0$ .

**Figure D.1**  
**Bias with and without Conditioning as a Function of Lead Rebate Coefficient—Baseline Model**



Notes. Based on 10,000 Monte Carlo simulations. Based on the selection equation (41), where there is a correlation between the probability of a False Negative and lagged expenditure.

upward bias in the MPC. When  $\zeta = 0$  there is no selection and the rebate coefficient is downward biased by  $-\$10$  due to classical measurement error.

When  $\zeta = -575$ , denoted by the dashed black line in Figure D.1, the upward bias is around  $\$180$ . The bias falls all the way to  $-\$10$  when we condition on lagged expenditure. Including the lagged rebate control in Table 3 in the paper reduces the contemporaneous rebate coefficient by  $\$85$  in the full sample and  $\$280$  in the rebate only sample. Our Monte Carlo estimates of the bias are right in the middle of these two values. Thus, our Monte Carlo suggests that selection on lagged expenditure, once disciplined by the lead rebate effect, can quantitatively explain the drop in our MPC estimates once we control for lagged expenditure.

#### **IVF Other Models of Misspecification and Selection**

We also explored various other data generating process where lagged expenditure is correlated with lagged rebate receipt. As an example, figure D.2 shows two of these alternative simulations. Model B shows an example of when there are both false negatives (as in our baseline model) and false positives where non-rebate recipients sometimes report receiving a rebate. Both the qualitative and quantitative results in model B are similar to model A.

Model C, explores a case where there is only false timing in rebate report. That is, every single rebate recipient eventually reports receiving a rebate, but the timing that they report receiving a rebate is correlated with lagged expenditure. In Model C, conditioning on lagged expenditure does not eliminate the upward bias, however, it does decrease it by around \$180.<sup>7</sup> A reduction similar to the fall in the rebate coefficient that we find in Table 3 of our paper after including lagged expenditure. We also simulated models with anticipation effects instead of selection on reporting. Anticipation effects lead to qualitatively similar results as shown here for Model C.<sup>8</sup>

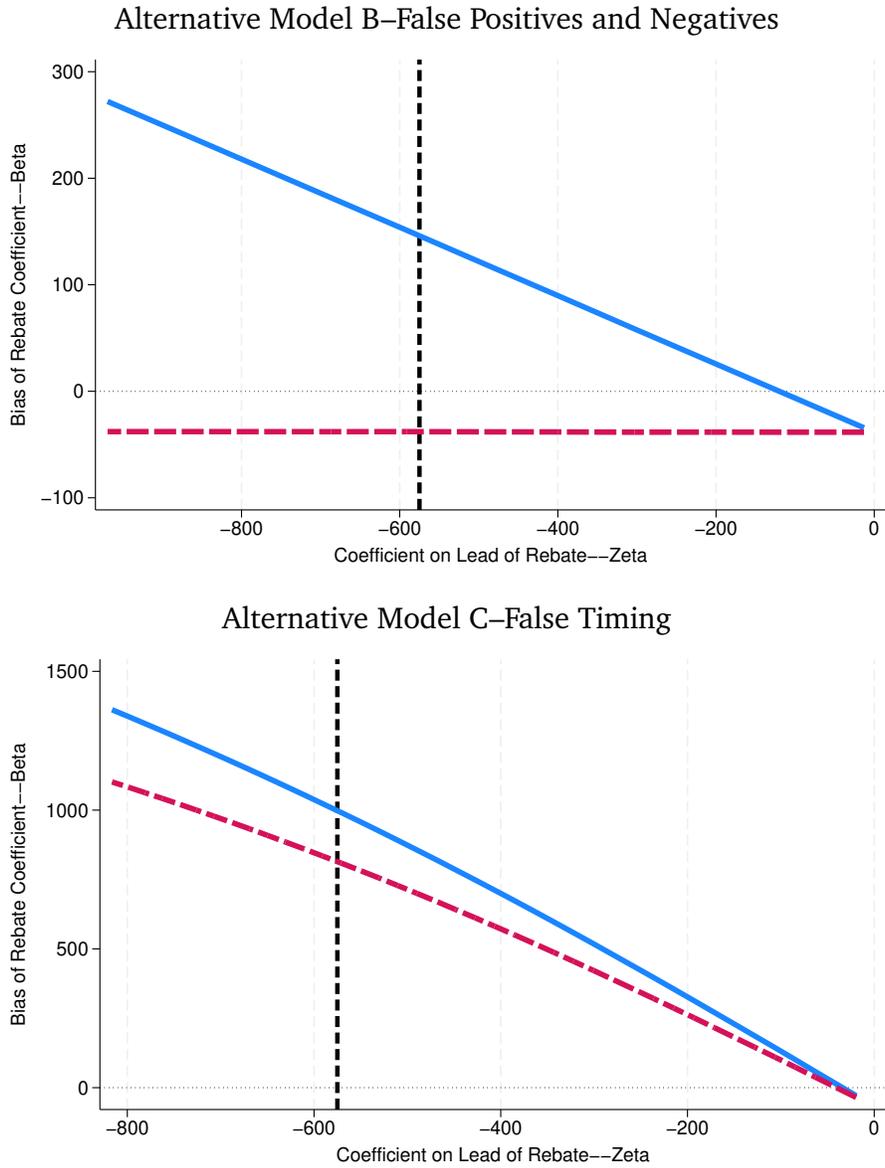
The three models disagree on how much bias remains after including the lagged expenditure control. When selection affects whether or not a rebate is reported, as in the baseline model and model B, the MPC estimates with the lagged control are very slightly downward biased. But in model C, when the timing is wrong, a substantial upward bias can remain. This suggests that our estimates are close to an upper bound for the true MPC.

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7. The lagged expenditure control in model C does not account for the mean reversion from the observations where a rebate is received, but is not reported. If a household misreports receiving a rebate at date  $t$  then the household will not report receiving at the households actual rebate date  $t^r$ . So while the false positives in this specification are conditional on lagged expenditure (and hence positive mean reversion), the false negatives are conditional only on the household having a false positive report for some observation. Following (41) a false positive means the observation has a high latent variable  $y_{it}^*$  and likely low lagged expenditure at time  $t - 1$ ; for some households time  $t - 1 = t^r$ , so false negatives are more likely for rebate recipients with low expenditure.

8. In the Monte Carlo Simulations with anticipation effects, the positive bias is eliminated when both lagged expenditure and the lead of the rebate are included as controls. However, in the actual data adding the lead of the rebate to our regression specifications does little to change the contemporaneous rebate coefficient, which provides suggestive evidence that anticipation effects are less important in our context.

**Figure D.2**  
**Bias with and without Conditioning as a Function of Lead Rebate**  
**Coefficient—Alternative Models**



Notes. Based on 10,000 Monte Carlo simulations. Model B is an alternative model that allows for both false negatives and false positives. Finally, in Model C, all rebate recipient report receiving their rebate at some point, but the timing of rebate report is correlated with lagged expenditure.

## E Details on the Additional Evidence for Nonrandom Rebate Reporting

### V.A Reported Rebate Date

Households in the CEX are surveyed every three months for a year in one of three interview schedules: the first month of the quarter (Jan, Apr, Jul, Oct), the second month (Feb-May-Aug-Nov), or the third (Mar-Jun-Sep-Dec). Table D.1 shows the interview schedules based on the month the household reports receiving the rebate. Panel A shows the entire recipient sample, while panel B shows only households that received a check rather than an Electronic Funds Transfer. In each case, the CEX interview schedule should not be related to the date of rebate receipt.

The first column of Panel A shows that in the overall CEX, there are an equal number of households in each interview group. Since the last two-digits of a household's SSN are effectively random, the households *actual* rebate date should have no correlation with the households interview schedule. However, households are more likely to report receiving the rebate the month prior to their interview. For example, households that report receiving their rebate in May are more likely to be interviewed in June. This suggests that some households may incorrectly recall the actual date of their rebate.

### V.B Expenditure by Rebate Schedule

Table D.2 shows that the households that report receiving a rebate the month prior to the interview also report higher spending.

The table is based on our baseline DiD specification, in which we now interact the rebate dummy with the number of months between when the household reports receiving the rebate and when they complete the interview:

$$C_{i,t} - C_{i,t-1} = \sum_s \beta_{0s} month_s + \beta'_1 \mathbf{X}_{i,t} \times + \sum_{s=1}^3 \beta_2^s I(ESP_{i,t}) \times I(\text{Interview } s \text{ months post rebate}) \\ + \sum_{s=1}^3 \beta_3^s I(ESP_{i,t-1}) \times I(\text{Interview } s + 3 \text{ months post rebate}) + u_{i,t}.$$

As an example, "Interview 1 month post rebate" means that the household reports receiving the rebate the month prior to the interview, while "interview 3 month post rebate" means the household reported receiving the rebate at the beginning of the 3-month recall period.

Households should be equally likely to show up in either the one, two, or three months post rebate bucket. Those that show up in the "Interview 1 month post rebate" will have had the least time to spend the rebate, while households in the "interview 3 month post rebate" bucket will have had the most time to spend their rebate. Despite

this, households in the “1 month post rebate” bucket generally have the largest estimated expenditure response. Conversely, households that report receiving the rebate three months ago have the smallest spending response. The differences are economically large, ranging from \$600 to over \$1000 dollars. And while all of these coefficients are noisy, we can reject that the coefficient on “1 month post rebate” is equal to the coefficient on “3 months post rebate” at the 15 percent level in all specifications.