OXFORD BULLETIN OF ECONOMICS AND STATISTICS, 85, 2 (2023) 0305-9049
doi: 10.1111/obes.12526

Identification Robust Empirical Evidence on the Open Economy IS-Curve

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Abstract

Existing empirical evidence on the Euler equation based on closed economy models suggests low responsiveness of aggregate consumption to changes in interest rates. We incorporate open economy features and consider extensions that include habit formation and hand-to-mouth consumers. For several open economies and applying econometric methods that are robust to weak instruments and structural changes, we continue to find low values for the elasticity of intertemporal substitution, implying a small effect of real interest rate changes on aggregate income. In some countries, structural changes are informative for identification, but otherwise aggregate data provide limited information to learn about IS-curve specifications.

I. Introduction

The New Keynesian framework has become a workhorse model for the analyses of monetary policy.[1] The model essentially reduces the economy to three major elements: a central bank that seeks to stabilize the output gap and to keep inflation as close to a target level as possible, a Phillips curve that expresses how a deviation of output from its potential level drives inflation dynamics, and an IS-curve that represents the intertemporal Euler equation. The IS-curve posits an inverse relationship between output and the real interest rate. Therefore, it provides an important channel for monetary policy to influence aggregate demand in standard dynamic stochastic general equilibrium (DSGE) models: consumers are assumed to substitute towards spending more when monetary policy lowers interest rates, and towards saving more when monetary policy brings interest rates up.

JEL Classification numbers: C1, C2, E1, F4.
We would like to thank the Editor Francesco Zanetti and three anonymous referees for their constructive feedback. We also thank Guido Ascari, Fabio Canova, Efrem Castelnuovo, Tim Robinson, Yong Song, Mark Weder and seminar participants at The University of Melbourne for very helpful discussions. The authors acknowledge generous support from the Australian Research Council under the grant DP170100697. Open access publishing facilitated by The University of Western Australia, as part of the Wiley - The University of Western Australia agreement via the Council of Australian University Librarians.

[1]See Goodfriend and King (1997), Clarida, Galí and Gertler (1999), Woodford (2003) and Galí, (2008) for exposition of the New Keynesian model.

The magnitude and direction of this effect are captured by the elasticity of intertemporal substitution (EIS), which is ‘‘a parameter of central importance in macroeconomics and finance’’, see Yogo (2004).

In this paper, we investigate open economy IS-curve models using limited-information generalized method of moments (GMM) that are robust to the problem of weak identification and parameter instability.[2] The main motivation of our work is to provide systematic evidence on the IS-curve model for open economies to various specifications of the model studied in the literature (and surveyed in section II) and report the findings in a unifying framework. We investigate extensions of the IS-curve based on the open economy framework of Clarida, Galí and Gertler (2002) and Galí and Monacelli (2005). These extensions include models with (*i*) habit formation in consumption as in Dennis (2009) and (*ii*) hand-to-mouth consumers (HTMC) as in Bilbiie and Straub (2012). The conventional IS-curve is purely forward-looking. The inclusion of habit formation in consumption, which is now a standard feature in macroeconomic DSGE models, for example, see Christiano, Eichenbaum and Evans (2005) and Smets and Wouters (2007), generates inertia in consumption dynamics, thereby lowering the responsiveness of consumption to interest rate changes. The introduction of HTMC also decreases this responsiveness since consumers who do not have access to financial markets simply consume their income, for example, see Campbell and Mankiw (1989), Bilbiie (2008) and Bilbiie and Straub (2012). All the open economy IS-curve models considered in this paper nest the closed economy as a special case.

Moreover, similar studies for the New Keynesian Phillips Curve (Kleibergen and Mavroeidis, 2009; Magnusson and Mavroeidis, 2014), the Taylor rule (Mavroeidis, 2010) and the (closed economy) Euler equation model (Yogo, 2004; Ascari, Magnusson and Mavroeidis, 2021) demonstrate the importance of employing econometric methods that are robust to the weak-instrument problem. Hence, we test the IS-curve specifications using methods proposed by Stock and Wright (2000), Kleibergen (2005) and Magnusson and Mavroeidis (2014) that are robust to weak identification. The method of Magnusson and Mavroeidis (2014) can also be interpreted as a parameter stability test robust to weak instruments and therefore can provide evidence on structural breaks in the IS-curve. We also apply a more recently developed method by Mikusheva (2021), which explores information from many potential instruments available for inference using a split-sample technique. We test the IS-curve specifications using these methods for a set of six countries fitting the features of open economies: Australia, Canada, New Zealand, Sweden, Switzerland and the United Kingdom. To the best of our knowledge, we are the first ones to empirically study open economy IS-curves using identification robust methods. Our findings read as follows.

First, the *aggregate* EIS, which measures the responsiveness of output to interest rate changes stemming from intertemporal substitution, is well identified and low in all countries. Confidence sets are mostly tight and include zero in all cases. This result is robust to weak identification and possible structural breaks. Several variants of both the closed and open economy models for the countries in our dataset point to low values of

[2]By focusing on limited-information methods, we refrain from imposing the entire DSGE structure that can lead to extraneous model misspecification.

the aggregate the EIS, thereby suggesting that small values of the EIS are empirical facts for most economies.

Second, the structural extensions to the baseline open economy model are weakly identified in most countries. Extending the baseline model to allow for habits, we find that higher degree of habits permits higher values of the individual EIS of the optimizers. However, this comes at a cost since the degree of habits is poorly identified, suggesting that instruments are weak. In models with HTMC, the fraction of HTMC is not well identified in Australia, Canada, New Zealand and the United Kingdom. As with habits, higher fraction of HTMC permits higher values of the individual EIS in these countries. For a given degree of openness, a limited aggregate response to changes in the real interest rate could arise either from low values of the individual EIS, a high degree of habits or a high fraction of HTMC. Therefore, in most countries, aggregate data have limited power to distinguish between alternative structural models. Only in Sweden and Switzerland, the fraction of HTMC is well identified and low; however, the EIS is also very low and not very different from the baseline model.

Third, there is evidence of instabilities in three countries in our dataset, namely Canada, Sweden and Switzerland. By exploiting information on instability within the sample, the stability test of Magnusson and Mavroeidis (2014) provides better identification of the degree of openness in these three countries, which are otherwise completely unidentified. However, for Australia, New Zealand and the United Kingdom, the degree of openness remains poorly identified. Moreover, for these latter three countries, we cannot reject the hypothesis that the structural parameters are stable over the entire sample, since the methods of Stock and Wright (2000) and Magnusson and Mavroeidis (2014) both provide very similar results.

The literature on the IS-curve has led to inconsistency between the values of the EIS that we expect from economic theory and its empirical estimates, giving rise to the so-called ‘‘IS puzzle’’. The EIS is often estimated to be not significantly different from zero (see, for example, Hall, 1988; Campbell and Mankiw, 1989), and when it is found to be significant, the EIS is often estimated to be small (or even negative, see Patterson and Pesaran, 1992; Goodhart and Hofmann, 2003; Gomes and Paz, 2013), suggesting that consumers are extremely risk-averse. Havránek (2015) conducted a meta-analysis on 169 published studies and concluded that the average estimate of the EIS in aggregate data is zero, once corrected for publication bias.[3] Two major causes of the IS puzzle suggested in the literature are: first, that the IS-curve is misspecified and cannot be fitted to empirical data (Stracca, 2010); and second, that the problem is purely an econometric one, arising from weak identification or time variance in the structural parameters in the model (Stock, Wright and Yogo, 2002).

Existing studies investigating the IS puzzle using identification robust methods either focus only on the conventional closed economy Euler equation model (Yogo, 2004) or consider model extensions within the closed economy framework but only focus on the US economy (Ascari *et al.,* 2021). We study several extensions of the Euler equation model using similar identification robust methods, but at the same time also incorporate an

[3]For studies estimating EIS using micro data, Havránek (2015) finds the EIS to be also very low and is around 0.3–0.4.

external sector. From a theoretical perspective, the assumption of an open economy is more realistic for the set of countries we consider. On the empirical side, weak-identification robust methods are more appropriate given identification problems with weak instruments. Nonetheless, we still find limited responsiveness of output to changes in the real interest rate for all countries in our dataset, suggesting that low values of the aggregate EIS are salient empirical facts.

The paper is organized as follows. The next section presents the open economy IS-curve specifications considered in this paper. Section III describes the empirical methodology and presents the data. Section IV presents the empirical results and the final section concludes the paper. In the online supplement, we provide detailed information on our dataset and several additional empirical results.

II. Open economy IS-curves

Baseline open economy model

Clarida *et al.* (2002) and Galí and Monacelli (2005) study a class of DSGE models with nominal rigidities to analyse optimal monetary policy in open economies. Following these authors, we incorporate an external sector capturing trade with the rest of the world. As in Clarida *et al.* (2002), there are two countries, home and foreign, that differ in size but are otherwise symmetric. The home country (H) has a mass of households $1 - \omega$, and the foreign country (F) has a mass ω. Otherwise, preferences and technologies are the same across countries. Within each country, households consume a domestically produced good and an imported good. Households in both countries also have access to a complete set of Arrow-Debreu securities that can be traded both domestically and internationally. We present the essential features of the model required to derive the IS-curve for an open economy framework and refer to Clarida *et al.* (2002) and Galí and Monacelli (2005) for a detailed analysis.

Let C_t be the following index of consumption of home (H) and foreign (F) goods:

$$C_t \equiv \left[(1-\omega)^{\frac{1}{\eta}} \left(C_{H,t}\right)^{\frac{\eta-1}{\eta}} + \omega^{\frac{1}{\eta}} \left(C_{F,t}\right)^{\frac{\eta-1}{\eta}} \right]^{\frac{\eta}{\eta-1}}, \tag{1}$$

where $\eta > 0$ measures the substitutability between domestic and foreign goods from the viewpoint of domestic consumers. The parameter $\omega \in [0, 1]$ is inversely related to the degree of home bias in preferences, and, therefore, can also be interpreted as a degree of openness. Let P_t be the corresponding consumption price index (that follows from cost minimization)

$$P_t \equiv \left[(1-\omega) \left(P_{H,t}\right)^{1-\eta} + \omega \left(P_{F,t}\right)^{1-\eta} \right]^{\frac{1}{1-\eta}}, \tag{2}$$

which, when log-linearized around a symmetric steady state satisfying the purchasing power parity (PPP) condition $P_{H,t} = P_{F,t}$, yields

$$\begin{aligned} p_t &\equiv (1-\omega)\, p_{H,t} + \omega p_{F,t} \\ &= p_{H,t} + \omega s_t, \end{aligned} \tag{3}$$

where lower-case letters denote the logs of the respective variables (with $p_t \equiv logP_t$) and $s_t \equiv p_{F,t} - p_{H,t}$ denotes the (log) terms of trade, that is, the price of foreign goods in terms of home goods.

The representative household in the home country maximizes

$$\mathbb{E}_t \left[\sum_{i=0}^{\infty} \beta^i \, u\left(C_{t+i}\right) \right], \tag{4}$$

where β is the subjective discount factor, and C_{t+i} is consumption in period $t+i$. We use the shorthand notation $\mathbb{E}_t[\,\cdot\,] = \mathbb{E}[\,\cdot\,|\mathcal{I}_t]$, where $\mathcal{I}_t$ is the set of information available to the consumer in period t. The instantaneous utility function, $u(\cdot)$, exhibits constant relative risk aversion (CRRA), that is

$$u\left(C_t\right) = \frac{C_t^{1-\frac{1}{\sigma}}}{1-\frac{1}{\sigma}}, \tag{5}$$

where σ is the elasticity of intertemporal substitution (EIS) or the inverse of the degree of risk aversion.

The first-order necessary conditions for consumption allocation and intertemporal optimization are standard:

$$C_{H,t} = (1-\omega) \left(\frac{P_{H,t}}{P_t} \right)^{-\eta} C_t, \tag{6}$$

$$C_{F,t} = \omega \left(\frac{P_{F,t}}{P_t} \right)^{-\eta} C_t, \tag{7}$$

$$Q_{t,t+1} = \beta \mathbb{E}_t \left[\left(\frac{C_{t+1}}{C_t} \right)^{-\sigma} \frac{P_t}{P_{t+1}} \right], \tag{8}$$

where $Q_{t,t+1} = 1/\left(1+i_t\right)$ is the price of the risk-free bond.

Log-linearizing equation (8) the above expression yields the baseline Euler equation:

$$\hat{c}_t = \mathbb{E}_t \hat{c}_{t+1} - \sigma \hat{r}_t, \tag{9}$$

where $\hat{c}_t$ denotes the log-deviation of consumption from steady state, and $\hat{r}_t = \hat{i}_t - \mathbb{E}_t \hat{\pi}_{t+1}$ denotes the log-deviation of the *ex-ante* real interest rate from the steady state. Henceforth, all variables with a hat denote log deviations from steady state.

A symmetric set of first-order conditions holds for citizens of the foreign country. In particular, given the international tradability of state-contingent securities, the intertemporal optimality condition can be written as

$$Q^*_{t,t+1} = \beta \mathbb{E}_t \left[\left(\frac{C^*_{t+1}}{C^*_t} \right)^{-\sigma} \left(\frac{P^*_t}{P^*_{t+1}} \right) \left(\frac{E_t}{E_{t+1}} \right) \right], \tag{10}$$

where an asterisk denotes the respective variable in the foreign country and E_t is the nominal exchange rate.

The law of one price, which implies $P_t = E_t P_t^*$ for all t, in conjunction with (8) and (10), and a suitable normalization of initial conditions, yields

$$C_t = C_t^*, \quad \text{for all } t. \tag{11}$$

Goods market clearing in the home and foreign countries implies

$$(1-\omega) Y_t = (1-\omega) C_{H,t} + \omega C_{H,t}^*, \tag{12}$$

$$\omega Y_t^* = (1-\omega) C_{F,t} + \omega C_{F,t}^*, \tag{13}$$

where Y_t and Y_t^* denote aggregate output in the home and foreign country respectively.

The demand curves for home and foreign goods by home citizens, (6) and (7), respectively, along with the analogues for foreign citizens and the law of one price imply that the consumer price index (CPI) based real exchange rate is unity. It then follows, after also taking into account (12) and (13), that

$$Y_t = \left(\frac{P_{H,t}}{P_t}\right)^{-\eta} C_t, \tag{14}$$

$$Y_t^* = \left(\frac{P_{H,t}^*}{P_t^*}\right)^{-\eta} C_t^*, \tag{15}$$

which, when log-linearized around a symmetric steady state, can be written as

$$\hat{y}_t = \hat{c}_t - \eta p_{H,t} + \eta p_t, \tag{16}$$

$$\hat{y}_t^* = \hat{c}_t^* - \eta p_{F,t}^* + \eta p_t^*. \tag{17}$$

Finally, combining (3) and (16) provides an aggregate demand schedule that relates domestic per capita output, per capita consumption and the terms of trade as follows:[4]

$$\hat{c}_t = \hat{y}_t - \eta\omega\hat{s}_t. \tag{18}$$

Substituting this expression into the Euler equation (9) results in our baseline open economy IS-curve specification[5]:

$$\mathbb{E}_t \Delta\hat{y}_{t+1} = \sigma\hat{r}_t + \eta\omega\mathbb{E}_t\Delta s_{t+1}. \tag{19}$$

[4]Since $P_F = P_H$ (in a symmetric steady state satisfying the PPP condition), the (log) terms of trade in steady state is zero and so $s_t \equiv \hat{s}_t$.

[5]Of course this equation is only valid in an economy without capital, durable goods, investment or a government. However, such small-scale models are regularly used for monetary policy analysis, see, for example, Lubik and Schorfheide (2007).

This specification is the same as in Clarida *et al.* (2002) for a two-country open economy model when preferences take the Cobb–Douglas form, that is, $\eta = 1$. The IS-curve can also be expressed in slightly different forms. For instance, as Clarida (2014) discusses, with identical preferences in domestic and foreign economy, goods market equilibrium implies a relationship between $\hat{s}_t$, $\hat{y}_t$ and $\hat{y}_t^*$ given by $\eta\hat{s}_t = \hat{y}_t - \hat{y}_t^*$, which can be used to substitute out for $\hat{s}_t$ in the equation above. Alternatively, in DSGE models the IS equation often appears in terms of log-deviation from its flexible price equilibrium, as discussed in Clarida *et al.* (2002), Galı́ and Monacelli (2005) and Clarida (2014) in the context of optimal policy. Since our focus is on providing identification robust empirical evidence on the IS-curve, we do not model the supply side of the economy. Hence, we abstract from re-writing the IS-curve in terms of the output gap but, instead, estimate the regression model obtained from the baseline open economy model (19).

Notwithstanding, if $\omega = 0$, that is, when there are precisely zero imports in the domestic household's consumption bundle, then from (18) $\hat{c}_t = \hat{y}_t$ and so (19) collapses to the Euler equation (9). Both (9) and (19) depict a negative reaction of output to the real interest rate due to intertemporal substitution, the magnitude of which depends on the EIS σ. However, in the specification of the utility maximization problem faced by the representative consumer, σ is also the inverse of the coefficient of relative risk aversion in consumption. The implication of $\sigma \approx 0$, a result found in several empirical studies as discussed before, is that the representative consumer is near-infinitely risk-averse, a conclusion that seems to be at odds with reality. This finding could potentially be driven by model misspecification, which, in this case, could be due to the omission of $\eta\omega\Delta s_{t+1}$ in the regression analysis. If variations in terms of trade growth are negatively correlated with changes in real interest rates (as it is in the data), then not accounting for terms of trade fluctuations in the regression analysis may lead to a downward bias in the EIS estimate.

Habits in an open economy

Fuhrer and Rudebusch (2004) and Paradiso, Kumar and Rao (2013), among others, empirically question the purely forward-looking nature of the IS-curve and introduce backward-looking dynamics into the model. This introduction is justified, for example, by allowing consumption habits: consumers consider past consumption choices when making current-period decisions. As argued by Fuhrer (2000), permitting habit formation makes the IS-curve a more realistic approximation of the relationship between consumption, interest rates and inflation.[6] Following Dennis (2009), we assume additive habits, so the utility function takes the following form:

$$u(C_t, H_t) = \frac{(C_t - H_t)^{1-\frac{1}{\sigma}}}{1-\frac{1}{\sigma}}, \tag{20}$$

[6]Some additional studies that use habit formation include Rees, Smith and Hall (2016) for Australia, Murchison and Rennison (2006) for Canada, Funke, Kirkby and Mihaylovski (2018) for New Zealand, Adolfson *et al.* (2008) for Sweden, Rudolf and Zurlinden (2014) for Switzerland, and Faccini, Millard and Zanetti (2013) for the United Kingdom.

and the stock of habits evolve as:

$$H_t = \gamma \left(C_{t-1}^{D} \tilde{C}_{t-1}^{1-D} \right), \tag{21}$$

where $\tilde{C}_t$ denotes aggregate consumption. The parameter $\gamma \in [0, 1]$ measures the degree of dependence on habits: if $\gamma = 0$ habits play no role and the model collapses to the baseline case, and if $\gamma = 1$ consumption is perfectly predetermined. The parameter $D \in \{0, 1\}$ is a dummy that determines the nature of the habits. If $D = 0$ habits are external, that is, the consumer is concerned with the level of her current consumption relative to the aggregate consumption in the previous period. If $D = 1$ then the consumer is concerned with the level of her current consumption relative to her own consumption in the previous period, therefore habits are internal.

Following Dennis (2009), the log-linearized Euler equation in case of external habits can be written as

$$\mathbb{E}_t \Delta \hat{c}_{t+1} = \gamma \Delta \hat{c}_t + \sigma (1 - \gamma) \hat{r}_t. \tag{22}$$

As seen from the above equation, habits modify the baseline closed economy model by introducing a lagged term $\Delta \hat{c}_t$, thereby changing the relative degree of backward-lookingness and forward-lookingness in the Euler equation. Moreover, external habits introduce a wedge between the individual EIS, σ, and the aggregate EIS, $\sigma (1 - \gamma)$. The case for internal habits is reported in the Supplement.

We extend the open economy IS-curve by allowing for habit formation in consumption as above. Similar to the baseline open economy model, equation (18) provides an explicit relationship between the representative domestic consumer's allocation of their income and the effective terms of trade facing the domestic economy. This result is easily substituted into the external habit specification in equation (22), yielding

$$\mathbb{E}_t \Delta \hat{y}_{t+1} = \gamma \Delta \hat{y}_t + \sigma (1 - \gamma) \hat{r}_t + \eta \omega \left[\mathbb{E}_t \Delta \hat{s}_{t+1} - \gamma \Delta \hat{s}_t \right]. \tag{23}$$

Extending the external habit case to an open economy introduces backward- and forward-looking dependence on the effective terms of trade, implying that habitual consumers consider the relative prices of imports and domestic goods, as well as the *ex-ante* real interest rate, when making consumption decisions.[7] When the economy is closed ($\omega = 0$), the model collapses to its respective habit counterpart, (22), and when the consumer displays no habitual behaviour ($\gamma = 0$), the model collapses to the open economy model in (19).

HTMC in an open economy

The assumption underlying the baseline Euler equation model is that all agents in the economy optimize their expected lifetime utility by substituting present consumption for saving, and vice versa. We relax this assumption by allowing for a certain proportion of the

[7] The open economy model with internal habits is discussed in the Supsplement.

population in the domestic economy that have no assets and simply consume their entire labour income in any given period, hereafter called hand-to-mouth consumers (HTMC).[8]

Following Bilbiie and Straub (2012), we let $\tilde{\lambda}$ be the proportion of HTMC in the economy, where $\tilde{\lambda} \in [0, 1]$. Hence, aggregate consumption is split between Ricardian (optimizing) households with access to saving (R) and non-Ricardian HTMC with no such access (N) as follows:

$$C_t = \left(1 - \tilde{\lambda}\right) C_t^R + \tilde{\lambda} C_t^N. \tag{24}$$

Log-linearizing this expression yields $\hat{c}_t = (1 - \lambda)\, \hat{c}_t^R + \lambda \hat{c}_t^N$, where $\lambda \equiv \tilde{\lambda} C^N / C$ is the fraction of total consumption consumed by HTMC in the steady state. The first-order conditions for consumption allocation and intertemporal optimization for Ricardian households are

$$C_{H,t}^R = (1 - \omega) \left(\frac{P_{H,t}}{P_t}\right)^{-\eta} C_t^R, \tag{25}$$

$$C_{F,t}^R = \omega \left(\frac{P_{F,t}}{P_t}\right)^{-\eta} C_t^R, \qquad \text{and} \tag{26}$$

$$Q_{t,t+1} = \beta \mathbb{E}_t \left[\left(\frac{C_{t+1}^R}{C_t^R}\right)^{-\sigma} \frac{P_t}{P_{t+1}} \right], \tag{27}$$

where the Euler equation (27) applies to Ricardian households only. Likewise, the first-order conditions for consumption allocation for non-Ricardian households are

$$C_{H,t}^N = (1 - \omega) \left(\frac{P_{H,t}}{P_t}\right)^{-\eta} C_t^N, \quad \text{and} \tag{28}$$

$$C_{F,t}^N = \omega \left(\frac{P_{F,t}}{P_t}\right)^{-\eta} C_t^N. \tag{29}$$

Under the assumption of a complete securities market, first-order conditions analogous to (25)–(27) must hold for the representative household in the foreign country (where, following Boerma (2014), we assume there are no HTMC in the foreign economy). The law of one price together with the Euler equations for home (Ricardian) and foreign households yield

$$C_t^R = C_t^*, \quad \text{for all } t. \tag{30}$$

Equation (30) implies that only Ricardian households share risk internationally.

[8]See Boerma (2014) for a small open economy version of the Calvo-type staggered price setting model with limited asset market participation.

Goods market clearing in the home and foreign countries yield

$$(1-\omega)\,Y_t = (1-\omega)\,\lambda C_{H,t}^N + (1-\omega)\,(1-\lambda)\,C_{H,t}^R + \omega C_{H,t}^*, \quad \text{and} \tag{31}$$

$$\omega Y_t^* = (1-\omega)\,\lambda C_{F,t}^N + (1-\omega)\,(1-\lambda)\,C_{F,t}^R + \omega C_{F,t}^*. \tag{32}$$

Combining the demand functions for home goods by Ricardian and non-Ricardian households, (25) and (28) respectively, together with the goods market clearing condition for home country, (31), and the law of one price, imply the following aggregate demand schedule

$$Y_t = (1-\omega)\left(\frac{P_{H,t}}{P_t}\right)^{-\eta} C_t + \omega\left(\frac{P_{H,t}}{P_t}\right)^{\eta} C_t^R. \tag{33}$$

Log-linearizing the above condition around a symmetric steady state yields

$$\hat{y}_t = (1-\omega)\,\hat{c}_t + \omega\hat{c}_t^R + \eta\omega\hat{s}_t, \tag{34}$$

where we have substituted $p_t - p_{H,t} = \omega\hat{s}_t$. The above equation can be rearranged as

$$\hat{c}_t = \left(\frac{1}{1-\omega}\right)\hat{y}_t - \left(\frac{\omega}{1-\omega}\right)\hat{c}_t^R - \left(\frac{\eta\omega}{1-\omega}\right)\hat{s}_t. \tag{35}$$

We can then use the equilibrium condition (35) to the apportionment of consumption between Ricardian and HTMC:

$$\left(\frac{1}{1-\omega}\right)\hat{y}_t - \left(\frac{\omega}{1-\omega}\right)\hat{c}_{R,t} - \left(\frac{\eta\omega}{1-\omega}\right)\hat{s}_t = (1-\lambda)\,\hat{c}_t^R + \lambda\hat{c}_t^N. \tag{36}$$

Rearranging this expression in terms of $\hat{c}_t^R$ and substituting into the Euler equation for Ricardian households, that is, the log-linearized version of (27), yields the IS-curve with HTMC in an open economy:

$$\mathbb{E}_t\Delta\hat{y}_{t+1} = \lambda\,(1-\omega)\,\mathbb{E}_t\Delta\hat{c}_{t+1}^N + [(1-\lambda)\,(1-\omega)+\omega]\,\sigma\hat{r}_t + \eta\omega\mathbb{E}_t\Delta\hat{s}_{t+1}. \tag{37}$$

Since HTMC simply consume their labour income, it follows that $\hat{c}_t^N = \hat{l}_t^N + \hat{w}_t^N$, where $\hat{w}_t^N$ is the real wage and $\hat{l}_t^N$ is hours worked of HTMC. We make assumptions that enable us to substitute for $\hat{c}_t^N$ and thus obtain equations containing only observable aggregate variables. Following Campbell and Mankiw (1989) and Ascari *et al.* (2021), we assume that a constant fraction of total labour income goes to HTMC. Hence, since consumption and income are in log deviations, it follows that we can replace $\Delta\hat{c}_t^N$ with the change in aggregate labour income. This is what we do in the empirical work below.

As with the other open economy models, (37) collapses to its closed economy equivalent when $\omega = 0$, that is, when domestic agents do not consume any imports. Moreover, (37) collapses to the baseline open economy model in (19) when $\lambda = 0$, that is, when all agents are optimizers and there are no hand-to-mouth consumers.

III. Empirical analysis

Econometric methodology

We use the GMM framework proposed by Hansen and Singleton (1982), where unobserved expectations terms are replaced by their realizations, predetermined variables are used as instruments and orthogonality conditions are obtained by assuming that residuals are uncorrelated with variables that are predetermined at time t. Considering the baseline IS-curve (19) with $\eta = 1$, we define the following regression specification:

$$\Delta y_{t+1} = \kappa + \sigma r_t + \omega \Delta s_{t+1} + u_{t+1}, \quad (38)$$

where κ is a constant that captures steady-state values of the real interest rate and output growth, $\Delta y_{t+1} = (1 - L)\, y_{t+1} = (y_{t+1} - y_t)$, $r_t = i_t - \pi_{t+1}$ is the *ex-post* real interest rate, and u_{t+1} is a disturbance term. To make our estimation robust to the presence of unanticipated shocks, we assume that $\mathbb{E}_{t-1}(u_{t+1}) = 0$, that is, only predetermined variables at time t can be considered as potential instruments. This assumption allows us to obtain the moment condition

$$\mathbb{E}\left[Z_t' u_{t+1}(\theta, \kappa)\right] = \mathbb{E}\left[Z_t'(w_t b(\theta) - \kappa)\right] = 0, \quad (39)$$

where Z_t is a vector of instrumental variables, $w_t = [\Delta y_{t+1}, r_t, \Delta s_{t+1}]$ and $b(\theta) = [1, -\sigma, -\omega]'$. All moments obtained from the models presented earlier can be cast in terms of $Z_t'(w_t b(\theta) - \kappa)$ with suitable redefinitions of w_t and $b(\theta)$.

Euler equation models are known to suffer from problems arising from weak instruments, see, for example, Yogo (2004) and Olea and Pflueger (2013). Therefore, we apply the S-test developed by Stock and Wright (2000) to produce confidence sets that are robust to the potential presence of weak instruments. Testing the hypothesis that $\theta = \theta_0$ is equivalent to testing $b(\theta) = b(\theta_0)$. The S-statistic for this test is then

$$S(\theta_0) = \frac{1}{T}\left(\sum_{t=1}^{T} Z_t'(w_t b(\theta_0) - \hat{\kappa}(\theta_0))\right)' \left[\widehat{V}(\theta_0, \hat{\kappa}(\theta_0))\right]^{-1} \left(\sum_{t=1}^{T} Z_t'(w_t b(\theta_0) - \hat{\kappa}(\theta_0))\right), \quad (40)$$

where $\widehat{V}(\theta_0, \hat{\kappa}(\theta_0))$ is a heteroskedasticity and autocorrelation consistent (HAC) estimator of the asymptotic variance of $T^{1/2}\sum_{t=1}^{T} Z_t'(w_t b(\theta_0) - \hat{\kappa}(\theta_0))$, since $u_{t+1}(\theta, \kappa)$ might follow a moving average process of order 1. We use the Newey and West (1987) estimator with four lags and Bartlett kernel. Finally, $\hat{\kappa}(\theta_0)$ is the continuously updated estimator (CUE) of Hansen, Heaton and Yaron (1996) for the untested parameter κ, which is obtained by minimizing $\mathrm{S}(\theta, \kappa)$ when imposing that $\theta = \theta_0$.

The S test is sufficient for testing the validity of the population moment conditions; however, it is not robust to potential instability in the moments or in the structural parameters. Magnusson and Mavroeidis (2014) propose an extension of the S test, the quasi-local level S (qLL-S) test, that examines whether the moment conditions are stable for a given vector of the structural parameters, θ.[9] The qLL$-$S test is a combination of the S-test with a test for the stability of the moments along the sample (qLL-$\tilde{S}$). Therefore, it can be thought of as using subsample information as additional instruments. The qLL-S test statistic takes the following form under the assumption that $\theta = \theta_0$:

$$\text{qLL-}\ S(\theta_0) = \text{qLL-}\tilde{S}(\theta_0) + \frac{10}{11} S(\theta_0). \tag{41}$$

The rejection of the test indicates the presence of instability that can be induced by (i) unstable parameters, or (ii) time variation in other parts of the economy, for example, regime shifts in monetary policy. In the first case, the qLL-S test can be interpreted as a structural change test that is robust to weak identification, implying that a non-rejection indicates parameter stability. In the second case, the qLL-S test can have more power than the corresponding S test if the information coming from structural changes elsewhere in the economy is sufficiently strong. In either case, the qLL-S test is complementary to the S test.[10]

In overidentified models, the S test has power against misspecification, making it difficult to interpret small confidence sets, and is not asymptotically efficient under strong identification. In addition, Doko Tchatoka and Dufour (2008) and Guggenberger (2012) show that S-type statistics can explode even for a slight violation of the moment conditions arising due to violation of the exclusion restrictions. Therefore, empty/small S and qLL-S confidence sets may result from instrument endogeneity rather than breaks in the parameters. To address these shortcomings and to ensure that the rejection of the model (e.g. empty confidence sets) is not due to the violation of the exclusion restrictions, we also compute confidence sets using the JKLM method as in Kleibergen (2005), which results from imposing the exclusion restrictions under the null hypothesis on the structural parameters. In addition to the S, qLL-S and JKLM sets, we also compute Kleibergen's extension to GMM of the conditional likelihood ratio (CLR) test of Moreira (2003). In the interest of brevity and because the JKLM and CLR sets turn out to be similar to the baseline results reported below, the main text does not show the CLR and the JKLM sets that are instead reported in the Supplement.[11]

In addition, the orthogonality conditions $E_{t-1}(u_{t+1}) = 0$ imply that any predetermined variable can be used as an instrument. Therefore, the number of potential instruments is unbounded. However, the S and qLL-S tests may be unreliable if the number of instruments is large relative to the sample size. As a result, we keep the number of instruments small when we compute S and qLL-S sets. This may be inefficient if information is thinly spread over many instruments, or if the most informative instruments are excluded from the set of instruments that we use. Accordingly, we consider a number of additional external instruments. Moreover, to address the potential threat to inference due to endogenous

[9] Magnusson and Mavroeidis (2014) recommend the qLL$-$S over other proposed tests because it is the most effective one in several cases of instability, and it is the most powerful test in the presence of persistent time variation as described in Elliott and Müller (2006).
[10] See Magnusson and Mavroeidis (2014) for technical details, including how to compute critical values for the test.
[11] See Figures S.2–S.5.

selection of instruments, we use a split-sample S test. This is a straightforward extension of a method recently proposed by Mikusheva (2021) to obtain reliable inference in linear instrumental variable models with time series data and many potentially weak instruments. In the Supplement we present information about the computation of the CLR, JKLM and split-sample S tests.[12]

We invert the results of these statistical tests to estimate the confidence sets. A 90% confidence set is the collection of all $\theta_0 \in \Theta$ for which the null assumption $\theta = \theta_0$ is not rejected. In case of the baseline open economy model $\theta = (\sigma, \omega)$, and, in case of the habits and HTMC models, $\theta = (\sigma, \gamma, \omega)$ and $(\sigma, \lambda, \omega)$, respectively. The parameters γ and λ are by nature restricted to the unit interval, that is, $\gamma \in [0, 1]$ and $\lambda \in [0, 1]$. As we are interested in realistic and applicable values of the elasticity of intertemporal substitution, we search within the space given by $\sigma \in [0, 4]$. We restrict our grid search for ω between $[0, 0.4]$. Note that ω corresponds to the import-to-GDP ratio (Galí and Monacelli, 2005), which is less than 0.4 for all countries in our dataset. Finally, following Lubik and Schorfheide (2007), we calibrate η, the parameter that governs substitutability between domestic and foreign goods, to 1 in our baseline analysis and report sensitivity of our results to alternative calibration of η in the Supplement.[13]

Data

We use aggregate quarterly time series data for Australia, Canada, New Zealand, Sweden, Switzerland and the United Kingdom. These countries are heavily reliant on international trade such that international relative price movements are likely to affect aggregate demand. Data are also readily available for relatively longer periods for these countries, which is important for the performance of the empirical tests.

We use logs of seasonally adjusted real gross domestic product for y_t, which implies that the IS-curves are generalized to the whole aggregate demand and applied to aggregate output, as often done in small-scale macro models; see, for example, Fuhrer and Rudebusch (2004) and Lubik and Schorfheide (2007).

The nominal interest rate i_t is taken as a quarterly average of monthly observations of rates on 90-day treasury bills or a comparable 90-day interbank rate, which is in line with prevailing practice in the literature, which allows for a direct link to monetary policy. Inflation π_t is computed as the log-difference between current and past-period levels of the consumer price index (CPI). The *ex-post* real interest rate r_t is the difference between the nominal rate at time t and the inflation rate at time $t + 1$.

We use real total compensation of employees as a proxy for the consumption of HTMC, except for Sweden where we use real net disposable income due to data availability. Log-differences are used in the empirical analysis to make this substitution valid. Clearly not all employee compensation are paid to HTMC, but assuming that a fixed proportion is, we can treat changes in employee compensation as a valid proxy for $\Delta\hat{c}_t^N$.

Finally, for terms of trade, we use the ratio of export to import price index.

[12] See section S.3 in the Supplement.

[13] For the model with internal habits, for which results are in the Supplement, we further calibrate the subjective discount factor β to 0.99, which is broadly consistent with equivalent calibrations in the business cycle literature; this corresponds to approximately 4% annual risk-free return in the steady state.

TABLE 1
Sample periods

Country	Sample period
Australia (AUL)	1968Q1 to 2018Q4
Canada (CAN)	1961Q1 to 2018Q4
New Zealand (NZD)	1989Q1 to 2018Q4
Sweden (SWD)	1980Q1 to 2018Q4
Switzerland (SWT)	1990Q1 to 2018Q4
United Kingdom (UK)	1963Q1 to 2018Q4

Table 1 presents the sample periods for the dataset used in the estimation. For each country, we consider the longest period for which data are readily available for the set of observables used in the estimation.

Other external variables included in the set of instrumental variables Z_t are oil price inflation, OECD's composite leading indicator (CLI) and (financial) uncertainty measure VXO.

Data were collected from a variety of sources including the Federal Reserve Economic Data (FRED) database and national statistical authorities; where this was not possible, gaps were filled using the International Monetary Fund's International Financial Statistics (IFS) database. The Supplement provides further details about the dataset.

IV. Empirical results

This section first presents our baseline results. Then, we report results based on external instruments, followed by results obtained from the split-sample method proposed by Mikusheva (2021) that is robust to many weak instruments.

Main results

We report the 90% confidence sets for the structural parameters based on the S and qLL-S tests in Figures 1–3. The set of instruments consists of a constant and the second lag of Δy_t, $(i_{t-1} - \pi_t)$ and $\left(i^*_{t-1} - \pi^*_t\right)$, where i^*_t is the US Federal Funds Rate and π^*_t is the US GDP deflator inflation rate.[14] We also include the second lag of terms of trade growth Δs_t as an additional instrument.[15] The number of instruments is kept small as otherwise the S and qLL-S tests may become unreliable. Panel (*a*) in each figure reports the two-dimensional 90% S set for (σ, ω) in the baseline open economy model (19); panel (*b*) reports three-dimensional 90% confidence sets for (σ, ω, γ) in the open economy model with external habits (23); and panel (*c*) reports three-dimensional 90% confidence sets for $(\sigma, \omega, \lambda)$ in the open economy model with HTMC (37). Panels (*d*)-(*f*) show the corresponding confidence sets based on the qLL-S test.

[14] Yogo (2004) shows that lagging the instruments twice assures that instruments are exogenous even under conditional heteroskedasticity.

[15] The Supplement shows that our results remain largely robust when we exclude terms of trade growth from the instrument set. See Figures S.9–S.11.

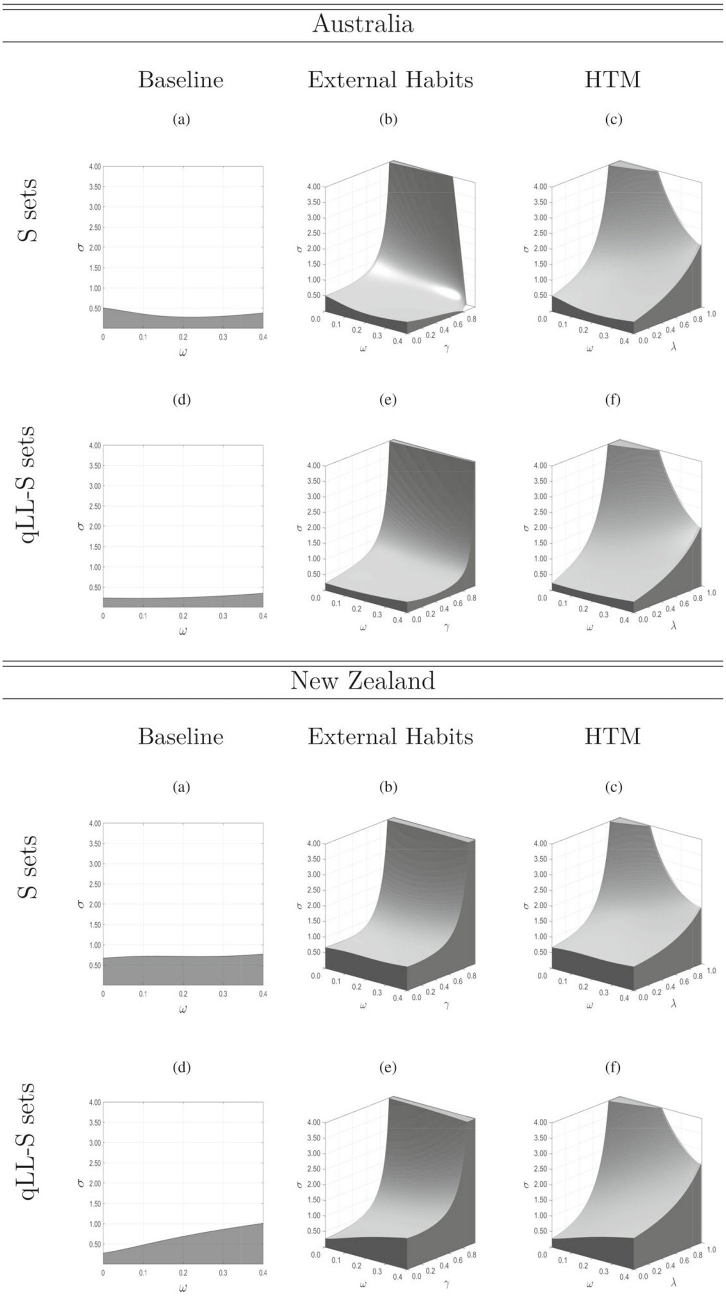

Figure 1. 90% S and qLL-S confidence sets for σ (individual EIS), γ (degree of habit formation), λ (fraction of HTM consumers) and ω (degree of openness) for Australia (1968q1–2018q4) and New Zealand (1989q1–2018q4). Instruments: constant, the second lag of Δy_t, $(i_{t-1} - \pi_t)$, $\left(i^*_{t-1} - \pi^*_t\right)$ and Δs_t. $\eta = 1$. Newey and West (1987) HAC with 4 lags

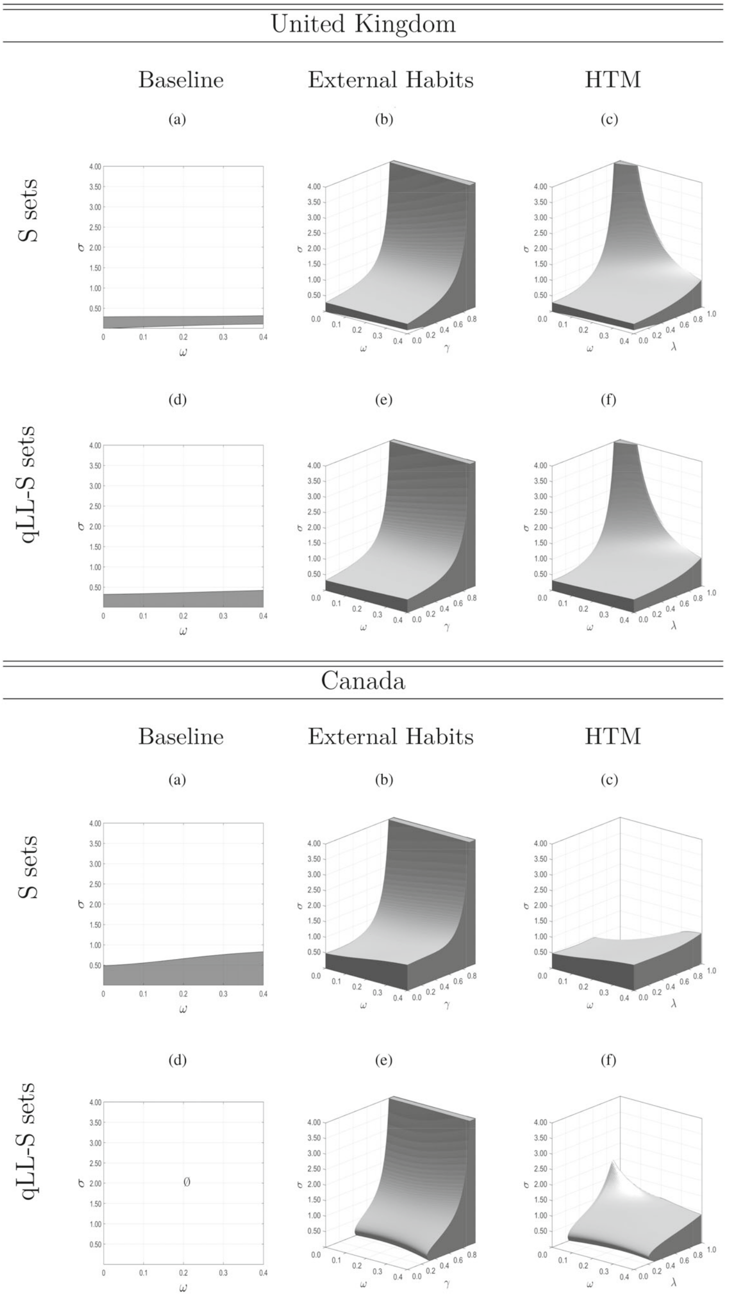

Figure 2. 90% S and qLL-S confidence sets for σ (individual EIS), γ (degree of habit formation), λ (fraction of HTM consumers) and ω (degree of openness) for United Kingdom (1963q1–2018q4) and Canada (1961q1–2018q4). Instruments: constant, the second lag of Δy_t, $(i_{t-1} - \pi_t)$, $\left(i^*_{t-1} - \pi^*_t\right)$ and Δs_t. $\eta = 1$. Newey and West (1987) HAC with 4 lags

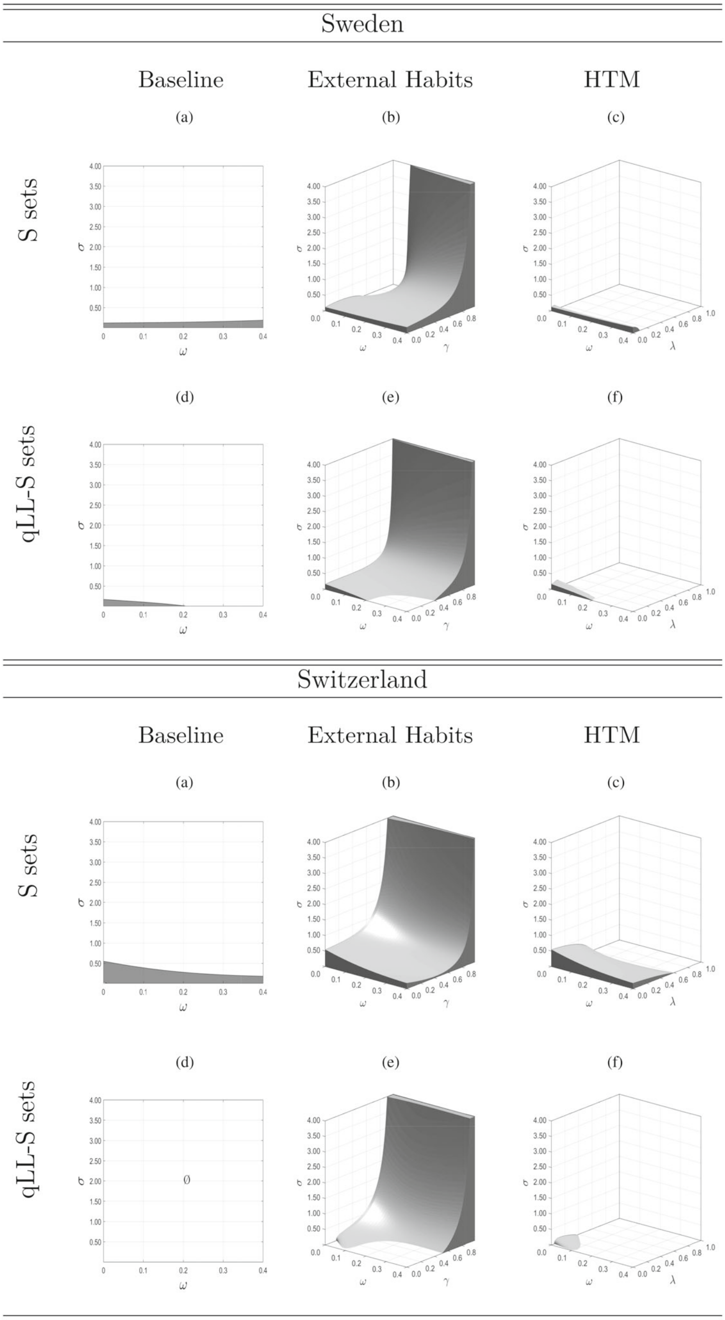

Figure 3. 90% S and qLL-S confidence sets for σ (individual EIS), γ (degree of habit formation), λ (fraction of HTM consumers) and ω (degree of openness) for Sweden (1980q1–2018q4) and Switzerland (1990q1–2018q4). Instruments: constant, the second lag of Δy_t, $(i_{t-1} - \pi_t)$, $\left(i^*_{t-1} - \pi^*_t\right)$ and Δs_t. $\eta = 1$. Newey and West (1987) HAC with 4 lags

First, we discuss the results for each of the models in turn, starting from the baseline open economy model. The 90% S sets in panel (a) in each figure show that σ is estimated to be lower than 1 in all countries. In fact, the confidence sets also include $\sigma = 0$. This is consistent with previous findings in the literature, for example, Yogo (2004), who use the same method and find that the EIS is small for several countries for the standard closed economy Euler equation model. In fact, this closed economy result can easily be seen from the y-intercept in panel (a) where $\omega = 0$. This suggests that small values of the EIS are not because of model misspecification due to omission of open economy features and that the IS puzzle remains even in open economy IS-curves. Additionally, we find that the degree of openness ω is completely unidentifed in all countries based on the S test. However, the $qLL - S$ sets in panel (d) show that sub-sample information turns out to be helpful for parameter identification in several countries where the qLL-S sets turn out to be smaller. For instance, ω turns out to be smaller than 0.2 for Sweden. Nevertheless, the admissible values of σ still remain small. Additionally, the 90% qLL-S sets for the baseline open economy model for Canada and Switzerland turn out to be empty, that is, there are no values of σ and ω for which the identifying restrictions of the model are acceptable at 10% significance level. Recall that the qLL-S test of Magnusson and Mavroeidis (2014) serves as a parameter stability test that is fully robust to weak instruments, and, therefore, this finding provides evidence of structural breaks in the baseline open economy IS-curve for Canada and Switzerland.

Turning to the model with external habit in panel (b), we can see that γ is poorly identified in all countries based on both the S and qLL-S tests. Note that higher values of σ become admissible as γ increases in all countries, but this comes at a cost since γ is left unidentified. In fact, the 90% confidence sets in most cases include $\gamma = 1$, at which point σ also becomes completely unidentified, that is, if $\gamma = 1$, then aggregate demand will not respond at all to changes in the real interest rate at any level of σ. In addition, as shown in the Supplement, the results are very similar when looking at models with internal habits.[16] This suggests that external and internal habits are empirically indistinguishable: they both fit the data but cannot be separately identified.

The results for the model with HTMC in the open economy are shown in panels (c) and (f) for the S and qLL-S tests, respectively. The fraction of HTMC λ remains poorly identified in all countries, except Sweden and Switzerland where λ is well identified and low, and σ is not significantly different from zero. In most other countries, higher values of σ become admissible as λ increases. In addition, the degree of openness ω remains poorly identified in most countries, with the exception of Sweden and Switzerland, where ω turns out to be smaller based on the qLL-S test.

Table 2 reports the fraction of the parameters from our grid space that are not rejected by the S and qLL-S tests, which correspond to the shaded areas in Figures 1–3. We also report this fraction based on the JKLM method for comparison.

We find that by exploiting information on the validity of the moment conditions over subsamples, the qLL-S test provides better identification of the baseline model for Canada, Sweden and Switzerland. In particular, the qLL-S sets are empty for Canada and Switzerland, therefore suggesting evidence of structural breaks in the baseline model.

[16]See Figure S.1.

TABLE 2

Fraction of the parameters that are not rejected at 10% significance level for the structural models

Country	Test	Baseline open	External habits	HTM
AUL	S	0.089	0.151	0.321
	$qLL - S$	0.068	0.182	0.310
	$JKLM$	0.119	0.157	0.300
NZD	S	0.182	0.337	0.330
	$qLL - S$	0.170	0.335	0.388
	$JKLM$	0.158	0.334	0.353
UK	S	0.059	0.275	0.175
	$qLL - S$	0.095	0.280	0.190
	$JKLM$	0.112	0.256	0.159
CAN	S	0.167	0.322	0.143
	$qLL - S$	0.000	0.231	0.130
	$JKLM$	0.212	0.349	0.203
SWD	S	0.040	0.145	0.003
	$qLL - S$	0.015	0.131	0.002
	$JKLM$	0.036	0.121	0.005
SWT	S	0.080	0.182	0.034
	$qLL - S$	0.000	0.179	0.002
	$JKLM$	0.071	0.168	0.037

This finding differs from the evidence on the Euler equation for the US economy in Ascari *et al.* (2021), where the qLL-S sets are bigger than their S counterparts. For the model extensions with habits or HTMC in these three countries, although the fractions in Table 2 are mostly similar, the shapes of the confidence sets in Figures 1–3 are somewhat different, suggesting that structural changes are informative for identification. Nevertheless, with the exception of the baseline open economy model for Canada and Switzerland, we cannot reject the hypothesis that the parameters of the model are stable over the entire sample, as otherwise the qLL-S sets would have been empty. In contrast, sub-sample information does not seem to provide sufficient additional identification for Australia, New Zealand and the United Kingdom, where the S and qLL-S sets are very similar.

To shed further light on the results, Figures 4–6 show the same confidence sets, but this time plotting the corresponding values for the aggregate EIS, φ, as opposed to the individual EIS, σ. For the baseline model (19), the individual and aggregate EIS are the same and therefore panels (*a*) and (*d*) in Figures 4–6 are the same as those in Figures 1–3. However, for the models with external habit or HTMC, there is a wedge between the individual and aggregate EIS. In case of the model with external habits (23), φ equals $\sigma(1-\gamma)$; while for the model with HTMC (37), φ equals $[(1-\lambda)(1-\omega)+\omega]\sigma$.

Figures 4–6 show that the estimates of the aggregate EIS φ are small across all models and countries. Table 3 depicts the range of values of φ that are not rejected at 10% significance level. Both the figures and the table show that aggregate EIS are more precisely estimated and include $\varphi = 0$ in almost all cases, which suggests that instruments are strong for the real interest rate. This can help us understand why the

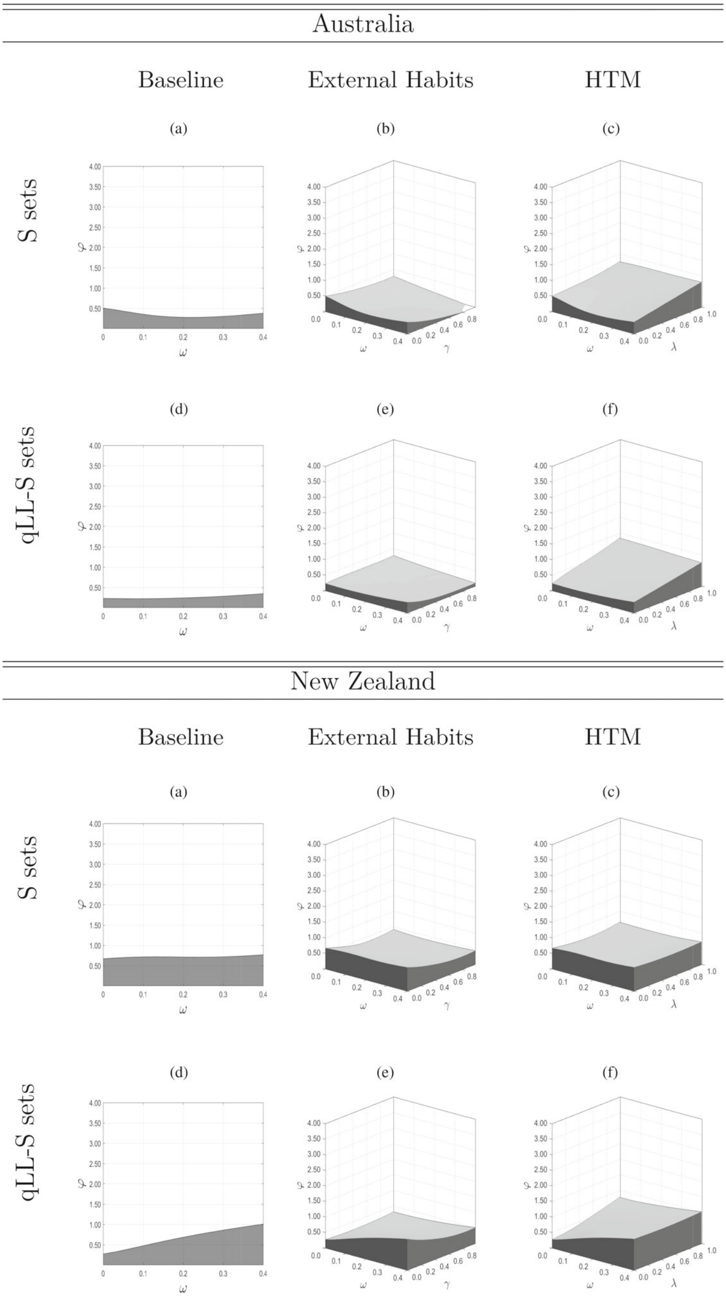

Figure 4. 90% S and qLL-S confidence sets for φ (aggregate EIS), γ (degree of habit formation), λ (fraction of HTM consumers) and ω (degree of openness) for Australia (1968q1–2018q4) and New Zealand (1989q1–2018q4). Instruments: constant, the second lag of Δy_t, $(i_{t-1} - \pi_t)$, $\left(i^*_{t-1} - \pi^*_t\right)$ and Δs_t. $\eta = 1$. Newey and West (1987) HAC with 4 lags

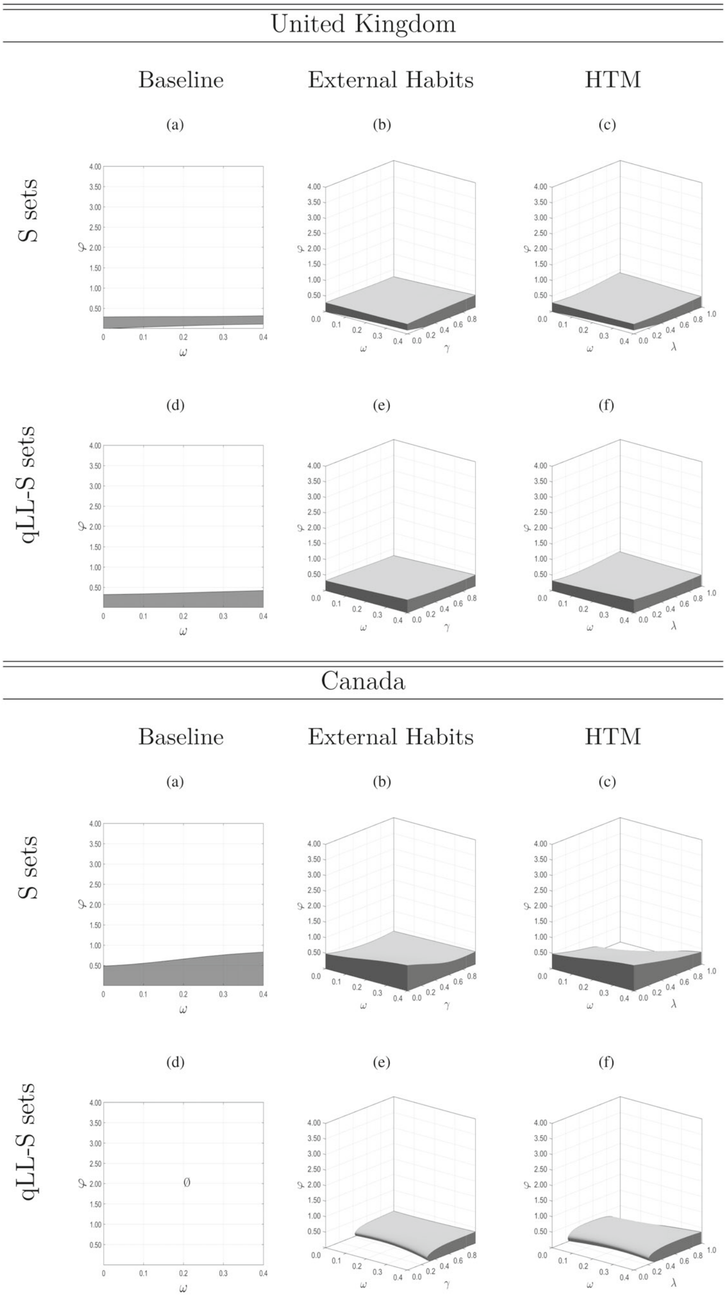

Figure 5. 90% S and qLL-S confidence sets for φ (aggregate EIS), γ (degree of habit formation), λ (fraction of HTM consumers) and ω (degree of openness) for United Kingdom (1963q1–2018q4) and Canada (1961q1–2018q4). Instruments: constant, the second lag of Δy_t, $(i_{t-1} - \pi_t)$, $\left(i^*_{t-1} - \pi^*_t\right)$ and Δs_t. $\eta = 1$. Newey and West (1987) HAC with 4 lags

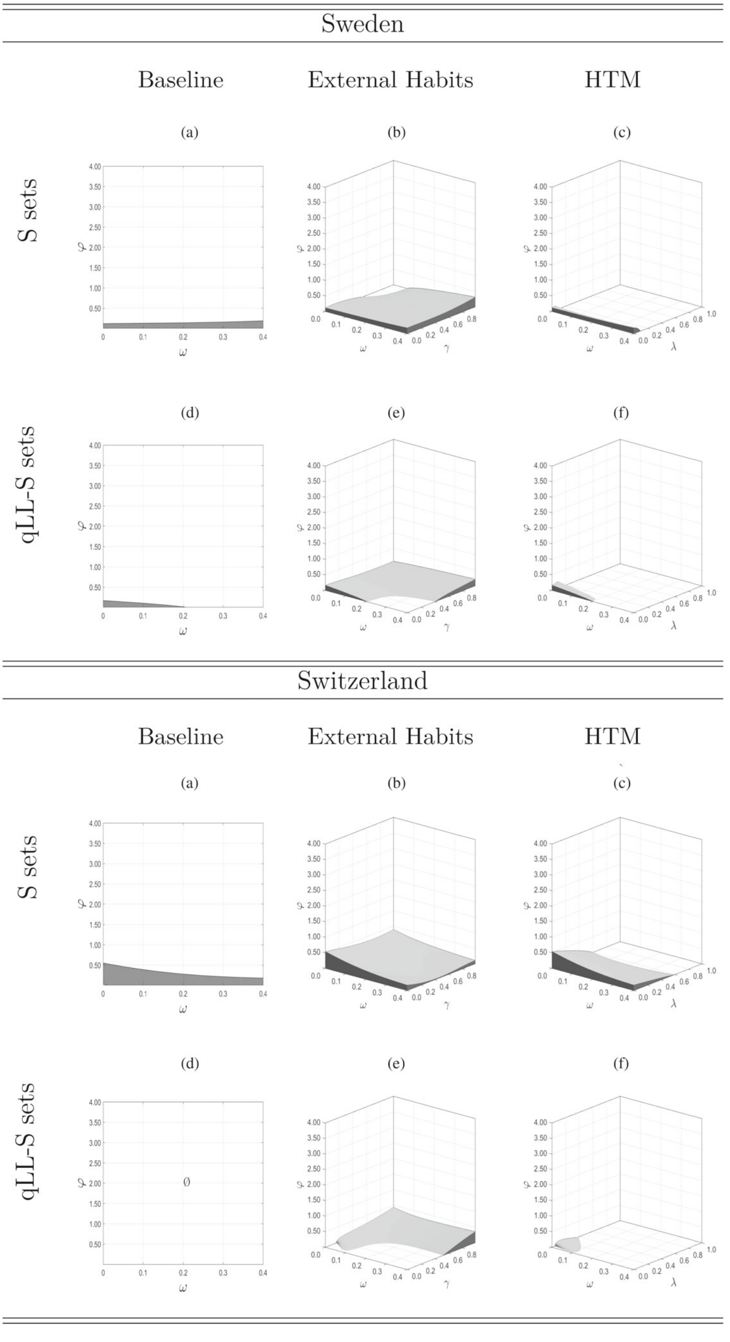

Figure 6. 90% S and qLL-S confidence sets for φ (aggregate EIS), γ (degree of habit formation), λ (fraction of HTM consumers) and ω (degree of openness) for Sweden (1980q1–2018q4) and Switzerland (1990q1–2018q4). Instruments: constant, the second lag of Δy_t, $(i_{t-1} - \pi_t)$, $\left(i^*_{t-1} - \pi^*_t\right)$ and Δs_t. $\eta = 1$. Newey and West (1987) HAC with 4 lags

TABLE 3

Range of values of aggregate EIS that are not rejected at 10% significance level

Country	Test	Baseline open	External habits	HTM
AUL	S	$[0-0.50]$	$[0-0.50]$	$[0-0.82]$
	$qLL-S$	$[0-0.34]$	$[0-0.34]$	$[0-0.84]$
	$JKLM$	$[0-0.79]$	$[0-0.66]$	$[0-0.74]$
NZD	S	$[0-0.77]$	$[0-0.76]$	$[0-0.76]$
	$qLL-S$	$[0-1.00]$	$[0-1.00]$	$[0-1.04]$
	$JKLM$	$[0-1.55]$	$[0-1.20]$	$[0-1.32]$
UK	S	$[0-0.31]$	$[0-0.38]$	$[0-0.40]$
	$qLL-S$	$[0-0.41]$	$[0-0.40]$	$[0-0.40]$
	$JKLM$	$[0-0.50]$	$[0-0.34]$	$[0-0.48]$
CAN	S	$[0-0.82]$	$[0-0.82]$	$[0-0.82]$
	$qLL-S$	$\emptyset$	$[0-0.36]$	$[0-0.40]$
	$JKLM$	$[0-1.18]$	$[0-0.94]$	$[0-0.92]$
SWD	S	$[0-0.18]$	$[0-0.32]$	$[0-0.18]$
	$qLL-S$	$[0-0.16]$	$[0-0.20]$	$[0-0.20]$
	$JKLM$	$[0-0.14]$	$[0-0.28]$	$[0-0.32]$
SWT	S	$[0-0.55]$	$[0-0.54]$	$[0-0.54]$
	$qLL-S$	$\emptyset$	$[0-0.42]$	$[0-0.12]$
	$JKLM$	$[0-0.67]$	$[0-0.80]$	$[0-0.52]$

structural parameters are poorly identified in some cases. For instance, the baseline model can reconcile small values of the aggregate EIS φ only if the individual EIS (or the inverse of the degree of risk aversion) σ is also small. The other models could instead admit high values of σ and still imply small values of φ. For instance, both the external habits and HTMC models could imply small values of φ for high values of σ if γ or λ (for a given ω), respectively, is also high. Likewise, when λ is high, the model with HTMC could result in high values of σ if ω is low, yet imply small values of φ. Therefore, if instruments are weak, resulting in poor identification of γ, λ or ω respectively, the identification of σ also turns out to be poor, despite φ being well identified and low. Only in the case of Sweden and Switzerland, where λ is well identified and turns out to be low, small values of φ also result in small values of σ.

Additionally, based on our results for the model extensions, note that the confidence sets for the parameters corresponding to habits γ or HTMC λ include zero values, with the exception of Canada and Switzerland. This suggests that for these other countries, even the baseline specification with low values of σ is consistent with the data and therefore neither habits nor HTMC are necessary to fit the data. In other words, with the exception of Canada and Switzerland, the confidence sets for the baseline model are all non-empty, suggesting that the identifying restrictions for this simple model are statistically acceptable at 10% significance level.

Lubik and Schorfheide (2007) estimate the structural small open economy NK model of Galí and Monacelli (2005) using full-information Bayesian estimation techniques for Australia, Canada, New Zealand and the United Kingdom, and also find that the value of σ

is mostly smaller than 0.5. Overall, the economic lesson to take away is that the aggregate EIS is small for all countries in our sample, implying a small direct effect of a change in the real interest rate on output stemming from intertemporal substitution. Moreover, aggregate data have limited power to distinguish between alternative theoretical models.

External instruments

Given the previous results, we explore a more extensive set of information contained in external instruments (external to the model that is) to ensure that the results are robust to instrument choice. These external instruments include (i) changes in the (log) oil price, (ii) OECD's composite leading indicator (CLI) and (iii) (demeaned and standardized) S&P100 Volatility Index (VXO). Oil price is measured using the West Texas Intermediate (WTI) spot crude oil price and can possibly pick up international conditions. The CLI is designed to provide early signals of turning points in business cycles showing fluctuations of economic activity around its long-term potential level. CLIs show short-term economic movements in qualitative rather than quantitative terms. VXO (one of the most popular measures of financial uncertainty) is used as a proxy for global uncertainty. Although VXO is a US-specific measure, we use it because of its availability over longer periods (available from 1962Q3). Caggiano and Castelnuovo (2021) show that US-based financial uncertainty strongly comoves with their constructed Global Financial Uncertainty (GFU) index, which, as with most other global uncertainty measures, is only available since 1990. Given the dominant role played by the US economy in world financial markets, this is perhaps not very surprising. Apart from oil price, the external instruments are in levels. As with lagged endogenous variables, we use the second lags as instruments.

To keep the number of instruments relatively small, we add one external instrument at a time to our baseline set of instruments. The resulting S and qLL-S confidence sets are reported in the Supplement.[17] In almost all cases, the confidence sets are very similar to the main results that use only lagged endogenous variables as instruments. The only exception is the baseline model for Switzerland, which turns out to be non-empty. However, the confidence sets remain very small with small admissible values of σ that include zero.[18] Therefore, additional information from external instruments does not seem to help much in identifying the structural parameters. On the upside, the robustness of the baseline confidence sets to using external instruments suggests that the results are most likely not driven by instrument choice.

Combining all instruments

The methods we used so far exploited information arising only from a few instruments at a time, even though the orthogonality conditions imply a large number of potential instruments (recall that any predetermined variable can be used as an instrument). This was done because the S and qLL-S tests become unreliable when the number of

[17]See Figures S.12–S.20.
[18]The confidence set for the baseline model for Canada also turns out to be non-empty when using all three external instruments at the same time but the admissible set remains small.

instrument is large relative to the sample size. However, use of many instruments could sharpen our inference if it happens to be the case that information is spread thinly over many instruments or if the most informative instruments are excluded from the set of instruments that we use. As a result, we now combine all the instruments that we have used so far in a single estimation. To address the potential threat to inference due to endogenous selection of the instruments, we construct a split-sample S set, which is robust to many weak instruments. We use approximately the first half of the sample to estimate the first-stage regression coefficients and the second half to compute the test statistic, following Mikusheva (2021) and as explained in the Supplement. In particular, we use a constant and the second lag of all the endogenous and exogenous variables as instruments. Figures S.21–S.26 in the Supplement show the results. For comparison, the upper two rows plot the S and qLL-S sets, respectively, while the bottom row plots the confidence set for the split-sample S set. Notable findings with respect to the split-sample S set are: (i) the baseline model for NZ turns out to be empty, (ii) confidence set for the baseline model for Sweden turns out to be somewhat larger accommodating values of σ greater than 1, and (iii) the baseline model for Switzerland admits higher values of σ but otherwise is not very well identified. Nevertheless, the results are overall in line with each other and with our baseline estimations for most countries in our dataset.

Further robustness

First, we compute the 95% confidence sets for the S and qLL-S tests. The confidence sets turn out to be marginally wider (as expected because of higher confidence level). Nevertheless, the results are very similar to those reported in Figures 1–3 and are relegated to the Supplement.[19] The only exception is the qLL-S confidence set for the baseline model for Switzerland, which turns out to be non-empty, but σ still remains small. The qLL-S set for the baseline model for Canada continues to be empty as before.

Next, we check the sensitivity of our results with respect to the calibration for the Armington elasticity η, that is, the elasticity of substitution between home and foreign goods. Following the calibration in Galí and Monacelli (2005) and Lubik and Schorfheide (2007), we previously set $\eta = 1$ in our analysis. However, a recent study by Imbs and Mejean (2015) shows that trade elasticity estimates decrease with the level of aggregation due to a heterogeneity bias. Using US data, the authors find evidence of higher values of macroeconomic trade elasticities when such elasticities are computed as a weighted average of sectoral elasticities. In light of this, we then set η to a higher value to check the robustness of our results. In particular, we now set $\eta = 3$. The confidence sets are reported in the Supplement.[20]

We find that in Australia, Canada and New Zealand σ turns out to be higher as ω increases in the baseline model. Nevertheless, with the exception of New Zealand, the qLL-S test suggests that σ is still mostly lower than 1. In New Zealand, higher values of σ come at a cost since identification of σ becomes weak as ω increases. For instance, when $\omega = 0.4$, the qLL-S confidence set for σ spans between 0 and 2. The confidence

[19] See Figures S.6–S.8.
[20] See Figures S.27–S.29.

sets for Sweden, Switzerland and the United Kingdom remain very similar. As before, the structural extensions including habits and HTMC are weakly identified in most countries, pointing to limitations of aggregate data in distinguishing between alternative structural models.

Finally, as discussed earlier, the qLL-S test can be thought of as using subsample information as additional instruments. It combines average information in the moment conditions over the sample with information on the validity of the moment conditions over subsamples. This subsample information can be relevant when there is time variation in other parts of the economy, such as monetary policy regime shifts. Central banks in most countries in our study adopted inflation targeting in the 1990s. To understand how this change in the conduct of monetary policy affects our results, we re-estimate the models using post-1990 data for Australia, the United Kingdom, Canada and Sweden.[21]

Figures S.30 and S.31 in the Supplement plot the 90% confidence sets. For each country, the upper panels show the qLL-S sets for the baseline sample period while the lower panels show the qLL-S sets for the post-1990 period. If structural changes around 1990 are important for identification, the qLL-S sets for the full sample will be smaller than the post-1990 sample. Indeed, we find that the qLL-S sets for the full sample are marginally smaller.[22] Nevertheless, the confidence sets are mostly similar, suggesting that the information coming from structural changes due to the adoption of inflation targeting is not strong enough to change inference on the parameters.

V. Conclusion

This paper investigates the empirical evidence on IS-curves for several open economies using aggregate data for Australia, Canada, New Zealand, Sweden, Switzerland and the United Kingdom. To overcome issues with identification, we use methods that are robust to weak instruments, parameter instability and structural changes. Several findings arise. First, we find that the aggregate EIS is well identified and low for almost all countries in our dataset. This finding is in line with existing empirical evidence based on closed economy models and therefore suggests that low values of the EIS are salient empirical facts for most economies. Second, extending the baseline model to allow for habits, we find that higher degree of habits permits higher values of the individual EIS of optimizing agents, but the habit parameter remains completely unidentified. Likewise, in models with HTMC, the fraction of HTMC is poorly identified in most countries with higher fraction of HTMC admitting higher values of the EIS. The only exceptions are Sweden and Switzerland, where the fraction of HTMC is well identified and low. Overall, our findings suggest that aggregate data have limited power in distinguishing between alternative theoretical models in most countries. Finally, we find that structural changes are informative for identification in some open economies, particularly for the degree of openness. Nevertheless, we still find limited responsiveness of output to changes in the interest rate, implying a flat IS-curve in most countries.

[21] Baseline results for New Zealand and Switzerland are based on data beginning in 1989Q1 and 1990Q1, respectively, and so we do not re-estimate them.

[22] The only exception is the HTM model for the United Kingdom for which the qLL-S set for the full sample turns out to be larger.

Final Manuscript Received: June 2021

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Supporting Information

Additional Supporting Information may be found in the online version of this article:

Appendix S1 Supporting Information