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Distribution and Gender Effects on the Path of Economic Growth: Comparative Evidence for Developed, Semi-Industrialized, and Low-Income Agricultural Economies

by

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ABSTRACT

This paper applies a robust empirical methodology, which considers issues relating to cross-country heterogeneity and cross-sectional dependence, to inspect the contributions of gender equality and factor income distribution to an economy's growth path. A dynamic model of aggregate demand is estimated on a unique panel dataset from 46 countries that are further grouped into developed (DC), semi-industrialized (SIEs), and low-income agricultural economies (LIAEs).

The empirical findings suggest that, overall, growth is driven by investment in the short run and domestic demand in the long run. In the short run, the results suggest that low female wages act as a stimulus to growth in SIEs but may promote contractionary pressures on demand in the long run. For LIAEs and DCs, the effect of improved labor market conditions for women—leaving men's constant—on demand-led growth conditions are positive in the short run but may harm long-term growth prospects.

In all, the empirical evidence, combined with the stylized facts about institutional and economic inequality, suggests that the impact of gender and income inequality on macroeconomic outcomes will differ depending on the economic structure and level of economic development.

KEYWORDS: Gender Equality; Demand-led Growth; Aggregate Demand; Functional Income Distribution; Economic Development; Autoregressive Distributed Lag Models (ARDL)

JEL CLASSIFICATION: E02; E12; E20; F41; I30; J03; J07

1. INTRODUCTION

This paper contributes to the literature by empirically examining the effects of the distribution of income by class and gender on aggregate demand (AD), and the potential implications for growth trends. We tie together both equity concerns, through the employment of structuralist macroeconomic models that emphasize the role of power relations between workers and capitalists in the determination of the distribution of income, and the rate of economic growth. In doing this, we are able to provide a pathway through which the persistent nature of gender inequity can be analyzed. Such a lens can then be used to emphasize the gendered power relations that exists at various levels of the economy and the implications they pose for income distribution and economic growth. This research then attempts to close the gap in the literature by employing empirical analyses to closely inspect the channels through which income and gender inequality affects growth.

We begin our analysis by highlighting the key features of gendered economic outcomes. The first is that women, on average, have less access to and control over material resources than men. Whether we measure this as wages, income, and wealth or as women's share of credit or landholdings, women on average fare worse than men. Second, a sharp gender division of labor persists, with women performing the largest share of unpaid labor (housework and care for children, the sick, and elderly). Within the sphere of paid employment, jobs are also gendered. Women are more frequently found in insecure, low-wage jobs while men are concentrated in higher income jobs with more security and benefits. While gender job segregation is everywhere apparent, it varies by country and especially by economic structure. The gendered distribution of jobs is a major factor influencing gender inequality in income and wages. Given this background, we employ a macroeconomic modeling strategy of demand-led growth, which highlights an underconsumptionist view of growth, to analyze the potential effect of income distribution and gender wage inequality on AD and growth. The contributions made are both theoretical and empirical.

On the theoretical side, this study attempts to engender the post-Kaleckian demand-led growth model by assuming a role for gender as a key determinant in the relevant categories of the behavioral components of AD. The Kaleckian approach, which emphasizes the role of monopoly power in determining the functional income distribution, is combined with a Marxist-leaning framework that attributes the distribution of wages between male and female workers to their relative bargaining power. This inclusion of a gender discriminatory effect (depicted by women's share of the wage bill) allows economies to simultaneously display wage-led and profit-led characteristics.

On the empirical side, our contribution is threefold. First, with the inclusion of a wage bill division between male and female workers, a conflictive or dual growth regime is possible. For example, a country can exhibit a female wage-led demand regime while simultaneously being profit-led with respect to the functional distribution of income. Secondly, a recent panel data approach that takes into account cross-country heterogeneity, dynamics, and the possibility of cross-sectional dependence is employed. Finally, to shed light on possible heterogeneity of the effects across countries at various stages of development, we also report separate results for high-income, middle-income, and low-income countries or, alternatively, advanced/developed economies (DC), semi-industrialized economies (SIEs), and low-income agricultural economies (LIAEs). This is done to account for the role that economic structure plays in influencing the relationship between gender equality and macroeconomic outcomes, and the nature of the persistence of this association. This categorization also enables us to empirically identify potential differences in the growth contributions of the gender and wage share variables across different levels of development.

Furthermore, attention is placed on the impact of a narrowing of the wage gap by raising women's wages (holding the average male wage constant). This approach is heterodox; unlike neoclassical theory, where wages are a product of labor's human capital endowments and marginal productivity, theoretical considerations following a heterodox approach allow for the influence of power structures in the determination of wages (Keynes 1936; Bowles and Gintis 1990; Figart, Mutari, and Power 2002).

These structuralist macroeconomic models, which emphasize the role of power relations between workers and capitalists in the determination of the distribution of income and the rate of economic growth, provide a pathway through which the persistent nature of gender inequity can be analyzed. Such a lens can then be used to examine the gendered power relations that exist at various levels of the economy and the implications posed for income distribution and economic growth. The objective of this paper is therefore to contribute to the analytical and empirical framework of the demand-led growth analysis by exploring the effects of changes in the functional income distribution and (gendered) personal income distribution on AD and its relevant components. Our model uses panel data on a sample of 31 countries over the period of 1970–2011.

The rest of the paper is grouped into sections. Section 2 presents the theoretical framework and section 3 introduces the data and methods of econometric analysis employed in the empirical investigation. In section 4, we report the empirical results. Section 5 derives and discusses the implications of the results; in that section, we also take a critical look at the examined relationships in the three income-grouped countries. Section 6 concludes and provides brief policy implications.

2. THEORETICAL MODEL

In macroeconomics, economic aggregates are usually assumed to be comprised of nongendered stocks and flows. However, disaggregation in macroeconomic modeling can provide for more efficient analyses and this includes disaggregation by gender. A proper rationale for incorporating gender into macroeconomic analyses is not one based on the physiological classification of economic agents as male or female but rather on the structural nature of persisting social strata and the effect that such relations impose on the behavior of economic models.

Demand-led growth models in the heterodox/Keynesian tradition have long emphasized the importance of income distribution and AD with relevance to wages as a source of demand to stimulate growth. Kaleckian macroeconomic models assume a significant role for the aggregate level of spending, which is a function of the distribution of income between workers and capitalists. This is a structural rather than an individualist disaggregation based on the different economic functions of workers and capitalists and corresponds to the institutional division between firms and households. This is a framework from which the employment of gender as an organizing framework in macroeconomics can be effectively introduced, as any such disaggregation ought to be based on a similar understanding of the way in which gender as a social institution impinges on or constrains the behavior of the macroeconomy.

2.1. Gender (Equality) Driven, Demand-Led Growth

The delineation between wage-driven and profit-driven growth is a crucial component of the neo-Kaleckian growth theory, and at the crux of this framework is the role of functional income distribution in determining a country's underlying growth regime. This conclusion is drawn from the logic that the distribution of national income between wages and profits has a considerable effect on aggregate saving and investment, which in turn affects the rate of capacity utilization and the rate of profit, respectively, thereby also affecting the capital accumulation rate. This theoretical framework does however ignore that several factors can cause distortions in the distribution of income and wealth across households.

While the functional income distribution is very important in explaining growth due to its effect on the rate of saving and capital accumulation, it is also important to account for the effects of the size distribution of income, which reflects how wages and profits are distributed within households.

We extend Bhaduri and Marglin's (1990) model of the original Keynes/Kalecki framework by incorporating a gendered distribution of wages and the impact of this on AD. We also allow wages to be weakly exogenous, firstly because we account for the role of wage rates in the determination of the price level while understanding that this codetermination of wages with the real level of output can be mitigated by policies allowing for fluctuations in the exchange rate in the case of an open economy—making changes in the real wage neither strictly endogenous or exogenous. This is especially true in open economies, as participation in international trade increases exposure to global price fluctuations while also fostering a push for competitiveness.

2.2. Aggregate Demand (AD) Components

Assuming an open economy with no government intervention, production can be assumed to occur through three different streams: stream one produces investment goods (I); stream two produces goods for capitalists and workers (C), and the third produces goods for foreign trade (NX). The public sector is captured by G, which is treated as exogenous. The behavioral function for AD can then be stated as:

$$Y=C+I+G+NX \tag{1}$$

where

$$C=f(Y,RW,WS,LP,DH) \tag{2}$$

$$I=f(Y,WS,INT,RW,DH,DB) \tag{3}$$

$$NX=f(Y,RW,WS,FY,EX) \tag{4}$$

where Y, RW, and WS are disposable income, the ratio of female to male wages, and the wage share, respectively. DH and DB are the household debt-to-GDP ratio and business debt. INT, FY, EX, and LP represent the long-term interest rate*,* foreign income in real terms*,* the nominal effective exchange rate, and female labor force participation, respectively.

Due to our focus on the private sector in this study, government expenditure (G) is treated as exogenous and as such does not feature in our empirical considerations. In the following subsections, we provide rationale for incorporating the main right-hand-side variables in the specification of the above-stated behavioral equations.

2.2.1 Consumption

Consumption is a component of AD and is assumed to be dependent on income. There are, however, other factors that affect the consumption behavior of individuals and households, such as changes in the interest rate, emulation effects, and the distribution of income.[1] Equation (2), above, extends the Keynesian aggregate consumption function, which is a function of autonomous income and disposable income (Y), to include the functional income distribution (WS) and personal income distribution by gender (RW). Y is expected to have a positive effect on aggregate consumption. WS is also expected to have a positive effect on consumption. This is assumed from the Kaleckian hypothesis that wage earners (workers) spend a greater proportion of their income than profit earners (capitalists). Therefore, an increase in the wage share should have a positive impact on aggregate consumption, as observed in Onaran and Galanis (2012) and Hein and Vogel (2007).

While a number of studies have examined the effect of the functional income distribution on aggregate consumption, only a few studies have looked into the effect of various other dimensions of inequality on aggregate consumption. This study aims to fill the gap in the literature by examining the effect of gender wage inequality on aggregate consumption, as wages are expected to influence an individual's average and marginal propensity to consume and this can vary between males and females.[2] The hypothesis is that women have a higher marginal propensity to consume (mpc) than men and, as such, an inverse relationship exists between gender inequality and aggregate consumption such that: $\frac{\partial C}{\partial RW} > 0$. We also expect that women's labor force participation (LP) may impact on their consumption expenditure either directly (because they are able to earn their own income by working) or indirectly

[1] Setterfield, Kim, and Rees (2016) posit that consumption can be driven by *emulation effects* when poorer households' consumption decisions are based on the desire to imitate the consumption patterns of more affluent households.

[2] See Keynes's analysis of consumption behavior as explained in *The General Theory* (Keynes 1936).

(because greater engagement of women in the labor market may imply that they are better able to organize in unions that advocate for higher wages for women relative to men).

Post-Keynesian models hypothesize that there is a dual effect of household debt on aggregate consumption (Dutt 2006; Palley 2010). While increased borrowing by households implies an increase in their disposable income, and therefore consumption, this also reflects an additional cost to these households as these loans must be serviced, putting downward pressure on consumption.

Research by various authors (e.g., Duesenberry 1949; Davanzati and Pacella 2010; Ryoo and Kim 2014; Setterfield, Kim, and Rees 2016: Setterfield and Kim 2016) hypothesizes that additional household borrowing has an initial effect of increasing consumption, as it is assumed that economic agents often try to emulate the consumption behavior of more affluent households, which could result in "consumption cascades." However, households (borrowers) need to service their debt with the implication that higher debt levels result in higher interest payments. Therefore, the overall effect of changes in households' indebtedness on consumption is ambiguous.

2.2.2 Investment

A common feature in the work of Keynes and Kalecki is the role attached to firms' rate of investment as a crucial determinant of output and employment. These investment decisions by firms are assumed to be influenced by their expectations of future demand, given their existing stock and their ability to finance such investment decisions through such means as procuring external debt or utilizing internal cash flow. This ability to finance investment independent of saving is an important theoretical precondition for both Keynes and Kalecki.

Some of these features influencing the investment decisions of firms—together with households' investment decisions—are summarized in equation (3), above. Aggregate investment is then expected to be composed of household and business investment. Y is expected to have a positive effect on I.[3] Similar to Stockhammer and Wildauer (2016), the effect of WS on I is ambiguous ($\frac{\partial I}{\partial ws} > 0 \; or \; \frac{\partial I}{\partial ws} < 0 \; or \; \frac{\partial I}{\partial ws} = 0$); while an increase in the wage share may indicate a loss of profits to capitalists or

[3] An increase in Y represents an increase in demand and the level of production, which, in turn, is likely to boost demand for capital and thus lead to greater investment.

businesses—which may discourage future investment—such a change in WS may promote residential investment (where workers are homeowners). It is expected that the effect of DH on investment in the long run is ambiguous because the accumulation of consumer debt results in a shift in the income distribution toward rentiers, who have a higher propensity to save (Asimakopulos 1983). In a similar vein, an increase in DB is expected to have an initial positive effect on investment unless businesses incur further debt to service their outstanding financial obligations. The long-term interest rate (INT) is expected to have a negative effect on investment.

The net effect of higher female wages on profits and thus investment may be positive or negative, depending on the economic environment. For example, the degree of firm mobility may determine the impact of higher female wages on investment.[4] In SIEs, women workers are concentrated in the export sector, which produces labor-intensive manufactured goods, business services, and nontraditional agricultural exports. In these regions, firms are often mobile, display monopsonist buyer of labor characteristics, and can easily trade globally through exports or export platforms—especially selling a bulk of their products in developed regions—leading to increased markups and greater comparative advantage (Busse and Spielmann 2003).

The implication then is that higher female wages can serve to dampen both investment and exports, producing an economic contraction and worsening the balance of payments (Bamber and Staritz 2016; Seguino 1997, 2007). In these countries, even if women's marginal propensity to consume is higher than men's (such that higher female wages stimulate consumption), the expansionary effect of higher female wages is unlikely to be large enough to offset the negative investment and export effects. Seguino (1997, 2007) explains that this phenomenon may be driven by the mobile nature of labor-intensive firms, which employ a large proportion of female workers in SIEs, where the export demand for the goods women produce is price elastic. This is especially so because there is often a large presence of these labor-intensive firms in such economies. The reverse may be the case for the effect of a more gender-equal wage setting in LIAEs.

[4] High firm mobility describes firms with low and limited sunk costs—including training costs—and easy firm entrance and exit. Mobile industries tend to be labor-intensive manufacturing firms as well as services (such as informatics, data processing, and possibly tourism) (Seguino 1997, 2007).

2.2.3 Net Exports

In the macroeconomic literature, changes in real foreign income (FY) and the nominal effective exchange rate (EX) are commonly assumed to be crucial determinants of a nation's export position. As in the case of investment, changes in the wage share (WS) represent redistribution from capitalists to workers, or vice versa, and, as such, affect production costs and, accordingly, exports. In addition, wages affect the price level in an economy, which further influences its international competitiveness. A priori, a negative association between WS and NX is hypothesized. We should note early on that to better represent our relationships between interest and the varying effect of gender and functional income distribution, separate import and export functions are estimated in this study. In the imports model, the potential dependence of export goods production on imported production inputs are accounted for by including exports as a right-hand-side variable in the import function.

The RW variable is used here to indicate not only the wage position of women relative to men but also their participation in the labor market—further represented by the inclusion of LP. While there is an obvious link between RW and NX using the same logic for the association between WS and NX, LP, which is expected to lead to higher wages in the long run due to increased influence and bargaining power in the labor market, also has significant effects on the export sector, as is evident in the literature (Seguino 1997; Berik 2000; Blecker and Seguino 2002). Ertürk and Darity (2000) highlight the dual effects of higher LP. Firstly, increased labor force participation by women often has the result of reducing both the time spent on unpaid caring labor—especially in the presence of rigid gender roles—and fertility rates, leading to increased earnings for women with potential negative impacts on the labor force, which, according to feminist theory, is a "produced means of production." Secondly, a lower wage position for women has a potentially positive impact on the composition and direction of production and thus exports, though which of these two cases has a higher effect on NX is expected to depend on the level of economic development (or structure) of a country.

As explained above, labor market conditions for women relative to men are expected to have different consequences in SIEs and LIAEs. For example, in LIAEs male labor is more concentrated in cash-crop production and nontradables (Collier, Edwards, and Roberts 1994; Seguino 2010a). Seguino and Were (2014) highlights that women, on the other hand, are more involved in subsistence agricultural production and sales; therefore, their involvement in this portion of production has no direct or immediate effect on international trade, as it does not directly influence global mobile investment, which is the opposite case for SIEs (as explained in the preceding subsection). An increase in women's wages (holding men's wages constant) will therefore be beneficial for domestic demand, at least in the short run.

Empirical research on intrahousehold bargaining suggests that men and women have different marginal propensities to consume, and these differences often depend on the type of consumption; for example, studies by Guyer (1988), Agarwal (1997), and Haddad (1999) find that men spend a larger proportion of their income on luxury goods rather than on basic household goods, while the opposite is true for women. This evidence appears to hold for countries at various levels of development. This could imply that for countries at a lower stage of development (e.g., LIAEs), consumption of luxury goods could be more import-intensive and where this is combined with a greater gender wage gap, household consumption on food, education, and healthcare may be lacking, which may be detrimental for short-run domestic demand and long-run productivity growth.

The long-run growth model will be more akin to a standard neoclassical supply-side model, in that the main drivers of long-run growth are technical progress, labor supply, and productivity growth.[5] For both SIEs and LIAEs, labor supply growth is positively correlated with increases in female incomes, as increased female labor force participation is correlated with less gender inequality (Blau and Kahn 2009; Seguino 2012). However, Kaleckian models describe the long run as a succession of many short-run periods and, as such, short-run effects on long-run growth or the quasi-equilibrium are expected.

2.2.4 Cumulative Effect on Income

The effect of consumption, investment, and net exports on aggregate income can then be analyzed by substituting equations (2) through (4) into equation (1). Following Stockhammer, Onaran, and Ederer (2008) and, more closely, Stockhammer and Wildauer (2016), the impact of a change in *WS* on *Y* can then be represented as:

$$\frac{dY}{dWS} = \frac{g_1}{1 - g_2} \tag{1}$$

[5] Here we refer to the inclusion of gender inequality indicators in a long-run productivity-driven growth model as is common in the mainstream literature, e.g., Dollar and Gatti (1999) and Klasen (2000). We, however, do not focus on this neoclassical model in this paper—especially the aspect of technology.

where

$$g_1 = \left(\frac{\partial C}{\partial WS} + \frac{\partial I}{\partial WS} + \frac{\partial NX}{\partial WS}\right) \text{ and } g_2 = \left(\frac{\partial C}{\partial Y} + \frac{\partial I}{\partial Y} + \frac{\partial NX}{\partial Y}\right)$$

where g_1 represents the initial short-term effect and can be defined as the resulting change in the level of AD due to a change in the functional income distribution level, $\frac{g_1}{1-g_2}$ represents the multiplier effect, and g_2 represents the marginal effects of Y on its components. Therefore, $g_1 > 0$ implies an initial positive impact of a higher WS on private demand; in this case, the economy is wage-led (and vice versa).

2.2.5 Gender Effects

Gender-specific discrimination against women in the labor market often manifests differently in developed and developing nations; for example, one expects that gender inequality in developed countries can be best identified through the differences between the wages of men and women, while in developing nations, differential access to wage employment may be a better indicator of gender-based discrimination in the labor market (Collier, Edwards, and Roberts,1994; Elson 1993; Cuberes and Teigner 2016). For this study, we are limited by data availability and, as such, specify the female–male wage ratio as the main gender inequality measure across our different country groups.

We then account for the cumulative gender effects, in terms of calculated elasticities, on the macroeconomic aggregates. Specifically, we expect that an increase in gender equality (described here as gender wage equality) that results in injections exceeding leakages (S +M < I +X) is expansionary (i.e., a redistribution stimulates AD), leading to an increase in output. A redistribution with this effect would be "gender cooperative," which refers to a redistributive effect of increased female wages on firms and the ratio of female to male wages, and otherwise "gender conflictive" when such an increase in the female–male wage ratio also results in a decline in male employment, thereby triggering an economic contraction (Seguino and Setterfield 2010).

3. DATA AND METHODS

In this section, the empirical relationship between personal and functional income distribution and AD is presented. Section 3.1 presents an overview of the data employed in the analysis; section 3.2. outlines the proposed empirical methodology.

3.1. Data

For this study, we compile an unbalanced panel dataset composed of 46 countries using annual data over the period 1985–2011. The 46 countries, presented in the appendix, are selected based on data availability and are collected from all geographical regions. This dataset is particularly large in comparison to the existing literature on the empirical determination of demand-led growth regimes.

For further comparative analysis, the panel is grouped into high-, middle-, and low-income countries, as defined based on the World Bank's classification, resulting in three panel datasets: 10 countries from 1985 to 2015 (panel 1) for low income, 12 countries from 1985 to 2015 (panel 2) for middle income, and 24 countries from 1985 to 2015 (panel 3) for high income.

Annual data for this study were derived from the Bank of International Settlements (BIS) and International Monetary Fund's (IMF) databases, the World Bank's World Development Indicators (WDI), Federal Reserve Economic Data (FRED), and the International Labour Organization's (ILO) data banks. A summary of the dataset for this macro panel analysis and the descriptive statistics is presented in the appendix, alongside the variations of data sourced for the gendered wage variables. Following the relatively long timespan (T = 31) of the dataset, one needs to consider several potential econometric issues often associated with panel data of this magnitude, namely, cross-sectional dependence (CSD), dynamics, slope heterogeneity, and the presence of unit roots.

3.2. Methods

To examine the nexus between income distribution and AD, a framework based on Bhaduri and Marglin's (1990) formulation of the Post-Keynesian demand-led growth model is proposed (Keynes 1936; Kalecki 1939, 1971). The adopted theoretical model for this analysis follows the approach by Stockhammer, Onaran, and Ederer (2008), and adds to this the gendered impact of income inequality within a panel framework.

We therefore provide a formal statistical analysis of distributional effects on AD following standard Keynesian representations as observed in equations (1) through (5), above. To shed light on possible heterogeneity of the effects across countries at various stages of development, reporting separate results for high-income, middle-income and low-income countries or, alternatively, advanced economies, SIEs, and LIAEs. This is done to account for the role that the economic structure plays in influencing the relationship between gender equality and macroeconomic outcomes, as well as the nature of persistence of this association.

Section 2.2 outlines the theoretical model for the empirical estimation, particularly, the determination of the multiplier effect and the implication of the numerator (private excess demand) and denominator terms (see equation [5]). To infer g_1, we calculate the sum of the marginal effects of WS on C, I, and NX. To proceed with our empirical analysis, one needs to consider the issues of nonstationarity, dynamics, heterogeneity, and CSD that may emerge from a macro panel dataset of this magnitude.

3.2.1 Preliminary Tests

Before testing for long-run cointegration between our variables of interest, we first check for the order of integration in our series using unit root tests. Specifically, we conduct panel data unit root tests allowing for homogenous and heterogenous slopes. We employ the Levin-Lin-Chu (LLC) unit root tests by Levin, Lin, and Chu (2002), which impose a homogeneity assumption, and the first-generation Im-Pesaran-Shin (IPS) test by Im, Pesaran, and Shin (2003) to allow for heterogenous autoregressive coefficients; both tests assume a standardized average of individual Augmented Dickey-Fuller (ADF) statistics to test the pooled null hypothesis of a unit root against a heterogenous alternative. We also employ a second-generation unit root test (the cross-sectionally augmented IPS [CIPS] of Pesaran [2007]) to additionally account for CSD among panel countries.

To test for a unit root, consider the conventional univariate ADF specification as follows:

$$\Delta y_{it} = \alpha_i + \beta_i t + (\rho_i - 1) y_{it-1} + \sum_{j=1}^{\rho} \delta_{ij} \Delta y_{it-j} + u_{it} \tag{2}$$

$$\text{with } H_{0:}\, \rho_i - 1 = 0, i = 1, \dots, N \tag{3}$$

$$H_{1:}\, \rho_i - 1 < 0, i = 1, \dots, N_1;\ \rho_i - 1 = 0, i = N_1 + 1, \dots, N$$

where y_{it} represents the respective panel series considered for country *i* at time *t*.

Under the LLC test, the null hypothesis is that $\rho_1 = \rho_2 = \cdots = \rho_N = 1$, while the alternative hypothesis assumes that $\rho_1 = \rho_2 = \cdots = \rho_N < 1$.

Where our panel series are I(0) or I(1), standard autoregressive distributed lag (ARDL) panel regressions can be relied on to produce efficient estimates, as demonstrated in a series of papers by Pesaran and others (see, e.g., Pesaran and Smith [1995], Pesaran [1997], and Pesaran, Shin, and Smith [1999]).

Table 1 presents the result of the LLC, IPS, and CIPS panel unit root tests. The various tests produce conflicting results on the unit root process. Using the LLC unit root tests, all variables other than imports, exchange rates, interest rates, and wage share possess a unit root. Results from the IPS test, which allows for heterogeneous coefficients across countries, show all variables other than female-to-male wage ratio, interest rates, and exchange rates to be nonstationary in levels. However, these variables become stationary upon first differencing. Using the simulated critical values of the CIPS, we find, for some variables, dissimilar results to the LLC and IPS tests for the levels. Upon differencing the relevant series, none of the variables were found to be > I(1).

Furthermore, given our panel, which consists of countries at different stages of development with differing social and economic conditions and, as such, conceivable structural differences, panel data methods that rely on pooling the data may result in potentially inconsistent and misleading estimations (Pesaran, Shin, and Smith 1999). It is therefore of importance that, in the presence of cross-sectional error dependencies and cross-country heterogeneity, an estimation strategy that adequately accounts for these features is employed. Disregarding such dependencies could lead to spurious inference and substantial bias in the estimated parameters of the specified model.

To this end, we carry out CSD tests under the null hypothesis of cross-sectional independence among countries, and poolability tests under the null of homogeneity of slopes across countries. To test for the existence of CSD among our countries, the general Pesaran (2004) cross-sectional dependence test

(CD_p) is employed, which under H_0 is distributed as a standard normal distribution ($CD_p \sim N(0,1)$ for $T_{ij} > 3$ and sufficiently large N, as is the case in this study.

The CD_p statistic can de defined as:

$$CD_p = \sqrt{\left(2T/(N(N-1))\right)}\left(\sum_{i=1}^{N-1}\sum_{j=i+1}^{N}\hat{\rho}_{ij}\right) \tag{4}$$

where $\hat{\rho}_{ij}$ is defined as the average pairwise correlation of the respective panel series with a null hypothesis of no CSD.

To test for slope homogeneity across countries, Pesaran and Yamagata's (2008) modified version of the Swamy (1970) poolability test ($\vec{\Delta}$ test) is employed and extended to the case of large N relative to T (as in this case) under homogeneity null under the condition of $(N,T) \rightarrow \infty$ without any restrictions on $\sqrt{N/T}$ with normally distributed error terms.

Following Pesaran and Yamagata's (2008) $\vec{\Delta}$ test for slope homogeneity in large panels, we first compute the standardized version of the Swamy (1970) poolability test given by:

$$\tilde{S} = \sum_{i=1}^{N}\left(\breve{\beta}_i - \acute{\beta}_{WFE}\right)' \frac{x_i' M_\tau x_i}{\acute{\sigma}_i^2}\left(\breve{\beta}_i - \acute{\beta}_{WFE}\right) \tag{5}$$

where $\breve{\beta}_i$ is the estimator from the pooled ordinary least squares (OLS), and $\acute{\beta}_{WFE}$ is the estimator from the weighted fixed effects pooled estimations of the regression models derived from $(\mathbf{1} \sim \mathbf{4})$ above; x_i is a $k \times 1$ vector of regressors; M_τ is an identity matrix; and $\acute{\sigma}_i^2$ is the mean square error from the OLS regression estimated for each cross-sectional unit.[6]

[6] Refer to Pesaran and Yamagata (2008) for a detailed presentation of the modified Swamy (1970) test and the broader definition of the estimators given in equation (10).

The standardized dispersion statistic $\left(\vec{\Delta}\right)$ is then defined as:

$$\vec{\Delta} = \sqrt{N}\left(\frac{N^{-1}\tilde{S} - k}{\sqrt{2k}}\right) \tag{6}$$

Under the null of homogeneity, for all *i* under the condition $(N, T) \to \infty$, when $\sqrt{N/T} \to \infty$ and with normally distributed error terms, the $\vec{\Delta}$ test has an asymptotic standard normal distribution.

Table 2 reports the results from the Pesaran (2004) cross-sectional dependence test (CD_p). The Pesaran and Yamagata (2008) standardized version of the Swamy (1970) slope homogeneity test $\left(\vec{\Delta}\right)$ results are reported in section 5. The cross-section dependence tests strongly indicate that the null hypothesis of independent cross-sections is rejected in favor of the alternative of dependent cross-sections for the analyzed panel series. Following the results for CSD, it is imperative that the presence of cross-sectional dependence be allowed for our empirical analysis.

Table 1: LLC, IPS, and CIPS Panel Unit Root Test Results

	Deterministic Trend	LLC	IPS	CIPS
A. Level				
Y	Trend, Intercept	1.855	13.902	-1.449**
C	Trend, Intercept	-1.085	-0.648	-1.129***
I	Intercept	4.798	8.295	3.946
X	Intercept	1.845	13.426	11.370
M	Trend, Intercept	-1.470*	0.711	-2.442**
WS	Trend, Intercept	-0.913***	-0.986	-0.570
RW	Trend, Intercept	-0.997	-2.130**	-3.556***
INT	Intercept	-16.682***	-8.285***	-0.963
EX	Intercept	-23.433***	-23.030***	-11.928***
LP	Trend, Intercept	-3.552***	-7.372***	-8.021**
DH	Intercept	1.738	11.836	1.092
DB	Trend, Intercept	4.679	9.002	-3.453***
FY	Trend, Intercept	-17.081***	-11.627***	-3.442*
B. First Difference				
ΔY	Trend, Intercept	-16.642***	-18.367***	-1.423***
ΔC	Intercept	-21.591***	-22.289***	-8.751***
ΔI	Intercept	-24.389***	-24.499***	-9.741***
ΔX	Intercept	-21.780***	-22.663***	-6.269***
ΔM	Intercept	-22.511***	-22.532***	-4.284***
ΔWS	Intercept	-14.979***	-18.650***	-1.275***
ΔRW	Intercept	-17.087***	-20.596***	-3.051**
ΔDH	Intercept	-22.098***	-11.534***	-2.109***
ΔDB	Intercept	-31.342***	-33.453***	-12.098***

Notes: We determine the optimal lag length using the Schwarz Information Criterion (SIC). Δ is the first difference operator. ***, **, and * denote rejection of H_0 at 1 percent, 5 percent, and 10 percent significance levels respectively.
The inclusion of a trend term is dependent on the observable characteristics of the series.

Table 2: Results of Cross-Sectional Dependence Test

	CD_p $Statistic$
Y	161.7164***
C	164.5973***
I	130.0207***
X	146.4400***
M	153.5564***
WS	28.31246***
RW	81.78625***
INT	94.31415***
EX	13.13568**
LP	33.65240***
DH	21.5856***
DB	86.9673***
FY	9.3910**

Note: ***, **, and * denote 1 percent, 5 percent, and 10 percent significance levels, respectively.

4. EMPIRICAL APPROACH

The empirical model is specified within a heterogeneous dynamic panel data setting based on a standard panel ARDL model (with p and q lags) with a multifactor error structure of the form:

$$y_{it} = \alpha_i + \sum_{j=0}^{q} \delta_{ij} x_{i,t-j} + \sum_{j=1}^{p} \lambda_{ij} y_{i,t-j} + u_{it} \tag{7}$$

$$u_{it} = \gamma_i' \boldsymbol{f}_t + \varepsilon_{it}, \tag{8}$$

where

$$x_{it} = a_i + b_i y_{i,t-1} + \Gamma' f_t + \nu_{it} \tag{9}$$

where i = 1,2, …, N; t = 1,2, …, T; $x_{i,t}$ is a $k \times 1$ vector of regressors for cross-section unit i at time t; δ_{ij} represents the $k \times 1$ coefficient vectors; λ_{ij} are scalars; and α_i and a_i represent a set of country-specific fixed effects capturing the impact of unobserved country-specific heterogeneity.

ε_{it} and ν_{it} represent the idiosyncratic errors that are assumed to be independently distributed across i and t with zero mean and constant variance; $\boldsymbol{f_t}$ is an $m \times 1$ vector of unobserved common factors that capture cross-sectional dependencies across countries; and γ_i' and Γ_i' are country-specific $m \times 1$ matrices of corresponding factor loadings.

This ARDL approach allows both the short-run dynamics and the long-run (quasi-)equilibrium effects of income distribution and gender equality on each of the variables in the macroeconomic and net export equilibrium conditions to be separately identified. Furthermore, the ARDL model specified in equation (8) takes into account certain properties of the data series, such as nonstationarity and unobservable common factors.

Equation (8) can be expressed as an equivalent error correction model (ECM) such that short-run dynamics are influenced by deviations from the equilibrium as denoted below:

$$\Delta y_{it} = \beta_i\left(y_{i,t-1} - \theta_i' x_{i,t-1}\right) + \sum_{j=0}^{q-1} \delta_{ij}'^{*} \Delta x_{i,t-j} + \sum_{j=1}^{p-1} \lambda_{ij}^{*} \Delta y_{i,t-j} + u_{it} \qquad (10)$$

Such that, $\beta_i = -\left(1 - \sum_{j=1}^{p} \lambda_{ij}\right)$; $\theta_i = {\sum_{j=0}^{q} \delta_{ij}} \Big/ {1 - \sum_{j=1}^{p} \lambda_{ij}}$; $\delta_{ij}'^{*} = -\sum_{m=j+1}^{q} \beta_{i,m}$;

$$\lambda_{it}^{*} = -\sum\nolimits_{m=j+1}^{p} \lambda_{i,m}$$

where β_i is the error correction coefficient that captures the speed of adjustment of deviations to long-run equilibrium; θ_i captures the long-run equilibrium relationship between our dependent and explanatory variables; and $\delta_{ij}'^{*}$ and λ_{it}^{*} define the short-run dynamics. A cointegrating relationship is inferred from a negative and significant β_i, which also provides an indication of the stability of the long-run cointegrating relationship.

This ECM specification is preferred to static or more restricted dynamic models because it accounts for heterogeneity across countries by allowing differing β_i and δ_i across the panel sample. It is also important to note that β_i is consistent whether the panel variables are I(0) or I(1) (Pesaran and Shin 1999). Also, given that these models are autoregressive, they do not suffer from endogeneity bias where sufficient lags are included (Chudik and Pesaran 20135).

Test results described in the previous section provide evidence of nonstationarity and CSD in our panel series. Therefore, the estimation of equation (12) for the theoretical models requires macro panel data techniques that allow for nonstationary series and CSD across countries while also providing estimates for all panel members. Another issue is that of endogeneity that arises from feedback between income distribution and AD in the specification of our theoretical models. In a seminal paper, Hansen and Tarp (2001) note that issues of endogeneity can be addressed to a reasonable measure with the use of dynamic panel estimation. By incorporating dynamics into equation (12), in this case as a lagged dependent variable, exogeneity restrictions may be somewhat relaxed. This technique allows for weakly exogenous regressors within a dynamic panel model, as in equation (15). Additionally, it allows for the possibility of feedback between variables in the respective consumption, investment, and net export equations; for example, past levels of the dependent variables can affect the regressors resulting in a case of weak exogeneity where such reverse causality exists.

Previously, we shed some light on the structural differences that may factor into the relationship between gender and the AD components, such as those between high-income/developed economies and low-income agricultural societies. These potential differences in gender behavior and outcomes from one country to another (depending on their level of development) may mean that the parameters of equations (1) through (4) are not equal for all panel countries. In such a case, panel regression techniques that require pooling of the panel series, such as the dynamic fixed effects (DFE) and generalized method of moments (GMM), may produce inconsistent and possibly misleading estimates. Following similar considerations, Pesaran and Smith (1995) develop a mean group (MG) approach, which allows for the separate estimation of the parameters for each cross-section; the mean of the estimated coefficients is then calculated to produce the reported coefficient estimates. Pesaran (2006) argues that error processes may suffer from cross-sectional dependence when panel cross sections are affected by factors not included in the estimation process or when there are spatial spillovers. Where such common factors are not adequately controlled for, misleading estimates may result.

We also employ the cross-sectional autoregressive distributed lag (CS-ARDL) model by Chudik and Pesaran (2015). As an additional robustness check, a reformulated ARDL specification (cross-sectional–distributed lag [CS-DL]) model is used to help avoid possible bias in long-run estimates resulting from inconsistency in the parameter of the lagged dependent variable. Both approaches allow for country-specific heterogeneity, error variances, and cross-country correlations; they are robust to endogeneity created by unobserved common factors, as well as omitted variable bias. The main advantage of the CS-DL regression is that it yields more precise long-run estimates than CS-ARDL when the time dimension of the data is not sufficiently long ($T<50$), as is the case in this study.

5. EMPIRICAL RESULTS

Using pooled time-series cross-section data (with N = 46 and T = 31, comprising 1,097 observations) on aggregate income (Y), aggregate private consumption (C), aggregate private investment (I), net exports (NX), adjusted wage share (WS), female-to-male wage ratio (RW), long-term interest rate (INT), and effective exchange rate (EX), consumption, investment, and net exports equations are estimated in a dynamic heterogeneous panel.

The analysis begins by choosing the optimal lag structure (*p,q*) for the ARDL, CS-ARDL, and CS-DL models using the Schwartz-Bayesian criterion (SBC). Chudik, Pesaran, and Yang (2016) explain the need for an appropriate lag length to ensure that the ARDL estimates are consistent, while also stating that where longer lags than necessary are employed, estimates with poor sample properties may result.[7]

Having already conducted formal tests to examine the properties of CSD and stationarity (unit roots) for our panel dataset and confirmed the prevalence of these issues, we proceed to estimate the heterogeneous dynamic ECM using the ARDL, CS-ARDL, and CS-DL estimators, results of which are reported in tables 3 through 12. It should be noted that while we report and discuss results for the ARDL, CS-ARDL, and CS-DL specifications, the analyses are based on results from the CS-ARDL model due to the possibility of endogeneity with some of our explanatory variables, and because of the presence of CSD. Moreover, Post-Keynesian macroeconomists consistently call attention to the role of the class distribution of income in influencing short-run outcomes, which has implications for the longer run. This feature of the discourse makes the CS-ARDL estimation especially useful in this study, as it provides the long-run estimates and short-run dynamics of the system.

Furthermore, the favorable results and relevant diagnostics (root mean square error [RMSE], *CSD* test statistic) obtained enable us to conclude the latter part of the demand effect estimations using results based on the CS-ARDL models augmented with three lags of cross-section averages.[8] In addition, due

[7] We also use the Akaike information criterion (AIC) as a cross-validation for the SBC method; we find that the most common lag length on the country-specific lag order is the same for both the AIC and SBC except in the case of the investment function. Eviews software is used in this model selection process.

[8] Chudik and Pesaran (2015) advise that the inclusion of sufficient cross-sectional lags (preferably three cross-sectional lags) is necessary to ensure allowance for the possibility of cross-sectional error correlations due to omitted common effects.

to the small sample bias the ARDL approaches face, the ECM is computed using a jack-knifed bias correction procedure on our estimates. This jack-knife procedure (originally proposed by Quenouille [1949] and further developed by Tukey [1956]) serves to correct for bias by a method of (re)sampling without replacement. This time series bias correction is carried out in Stata.

Moreover, we assume nonzero effects for our outcome variables. Evidence suggests that when it is negative and statistically significant and a long-run cointegrating relationship is established under a common correlated effects cointegration technique, it is impossible to distinguish a null effect from a very small effect (Muller 2004; Lane et al. 2015; Choi and Chudik 2019). Following this evidence, negligible effects, even when not statistically significant, are considered in estimating the cumulative distributional effects and in proposing gender and distributive effects.

For all the AD components, the baseline model is subjected to robustness checks. This check entails using the ARDL and CS-DL approaches as sensitivity checks for our CS-ARDL estimates. Furthermore, for all specifications, only the first lag short-run coefficients are reported; the short-run results are rarely statistically significant, and more so with the first lag results.

5.1. Consumption

Following the literature on demand-led growth, the baseline consumption function is estimated where C_t is a function of the income variable; Y_t is the variable capturing gender wage inequality (RW_t); and the wage share variable, WS_t, is based on a panel model of the general form[9]:

$$lnC_{it} = f(\, lnY_{it}, lnRW_{it}, lnWS_{it}) \qquad (11)$$

The ARDL, CS-ARDL, and CS-DL estimates of equation 16 are reported in table 3 using jack-knifed standard errors.[10] As earlier stated, the CS-ARDL estimator is preferred over others for this study due

[9]Assuming a linear relationship is a good starting point for our analysis, as most of our key explanatory variables (e.g., WS and RW) vary over only a narrow range of values. We then continue our analysis with a log-log specification as we expect that the impact of the independent variables on the dependent variables may in some cases be different from one level to the other. Furthermore, previous studies (Stockhammer, Onaran, and Ederer 2008; Onaran and Galanis 2012) also employ a log-log functional form given the interest in calculating marginal effects for the C, I, and NX models—which requires estimated elasticities. Finally, this transformation allows for easier interpretation of our estimates.

[10] All estimations for the AD components' equations are carried out using Stata 13.

to compelling evidence of cross-sectional dependence and the possible endogeneity issues with our variables.

5.1.1. Estimates of Long-run Effects

The least squares estimates obtained from the panel ARDL and distributed lag (DL) specifications in table 3a report the results for the baseline specification and table 3b shows the results for the extended regression when an additional explanatory variable (DH) is included. Each panel gives the error correction variant of the MG estimates of the long-run effects on aggregate consumption from changes in our explanatory variables.

With regards to the baseline specification, panel A of table 3a reports closely similar ARDL and CS-ARDL estimates for the long-run income effect on consumption, with highly significant elasticities varying from 0.63 to 0.99. These income elasticity results for the ARDL and CS-ARDL estimates appear more comparable under identical lag structures; when ρ=1 (i.e., ARDL [1,1] AND CS-ARDL [1,1]), the income elasticity of consumption ranges from 0.81 to 0.99. The long-run average coefficients on the wage share are positive under all model specifications, consistent with the demand-led growth literature (Stockhammer, Onaran, and Ederer 2008; Onaran and Galanis 2012; Hein and Vogel 2007). In addition, we observe that the wage share coefficients are statistically significant under all but the DL (1,1,1) model. Using our preferred CS-ARDL results, we expect that in the long run, a 1 percent increase in the wage share is associated with a 0.49 percent increase in aggregate consumption.

Table 3a: Regression Results for the Consumption Function (equation [2])

	ARDL		CS-ARDL		CS-DL
Lags	(1,1)	(2,2)	(1,1)	(2,2)	ρ =1
Regressions with key variables					
$\hat{x}_Y$	0.991*** (0.080)	0.938*** (0.009)	0.809*** (0.177)	0.633** (0.220)	0.529*** (0.059)
$\hat{x}_{RW}$	0.111** (0.166)	0.016 (0.031)	0.992** (0.080)	-0.856 (2.044)	0.048 (0.342)
$\hat{x}_{WS}$	0.331** (0.169)	0.210*** (0.031)	0.494** (0.036)	0.831* (0.223)	0.035 (0.265)
$\hat{\lambda}$	-0.413***	-0.174***	-0.549***	0.014	n.a.
CD	84.00***	92.49***	-0.50	-0.24	-0.69
$\vec{\Delta}$	16.857**				

Notes: $\hat{x}\_$ is an indicator of the regressor $x_{i,t}$ in equation (12).
$\hat{\lambda}$ is the speed of adjustment (ECT).
The CD test reports the CD statistics instead of p-values.
$\vec{\Delta}$ reports the Pesaran and Yamagata (2008) poolability test results for the estimated function.
Jack-knifed standard errors are reported inside parenthesis.
n.a. stands for not applicable when there are not enough observations to conduct the relevant estimations when using certain lags to deal with the potential serial correlation and CSD issues.
***, **, and * denote significance at the 1 percent, 5 percent, and 10 percent levels, respectively.

The coefficient of the female-to-male wage ratio is positive under all specifications (except for the CS-ARDL estimates when $\rho = 2$). It is worth noting that the RW coefficients are positive and significant only when one lag is imposed on the ARDL specification. The CS-ARDL and DL estimates, while mostly positive, are not significant and as such do not provide robust evidence to support the long-term relationship between aggregate consumption and gender wage equality. In this case, our ARDL (1,1,1) estimate suggests a 0.11 percent response of aggregate consumption to a 1 percent improvement in the proportion of the female wage bill.

Robustness to additional explanatory variables: Our baseline specification above imposes some simplifying assumptions on the consumption function. We therefore attempt to relax these assumptions by expanding the model estimated in table 4, above. Specifically, we consider the household debt-to-GDP ratio (DH) and the female labor force participation (LP) as additional factors that can potentially influence the aggregate level of consumption. We are constrained from adding more controls because of the reduced degrees of freedom.

We find the income elasticity of consumption estimates fall within a range of 0.72 to 1.03 when $\rho = 1$; all estimates are significant as in the baseline model, following a priori expectations. The ARDL, CS-ARDL, and DL long-run coefficients on the wage share are all positive. However, unlike the baseline results, only the CS-ARDL (1,1) estimates are statistically significant, with a coefficient of 1.168.

Table 3b: Regression Results for the Consumption Function

	ARDL		CS-ARDL		CS-DL
Lags	(1,1)	(2,2)	(1,1)	(2,2)	ρ=1
Regressions with additional variables					
$\hat{\boldsymbol{\theta}}_{\boldsymbol{Y}}$	0.994***	0.908***	1.027***	1.095***	0.717***
	(0.161)	(0.116)	(0.144)	(0.197)	(0.108)
$\hat{\boldsymbol{\theta}}_{\boldsymbol{RW}}$	0.792	0.362	0. 377**	-3.852	0.664
	(1.022)	(0.535)	(0.152)	(3.499)	(0.703)
$\hat{\boldsymbol{\theta}}_{\boldsymbol{WS}}$	0.329*	0.085	1.168*	4.348	0.121
	(0.403)	(0.643)	(0.674)	(3.803)	(0.115)
$\hat{\boldsymbol{\theta}}_{\boldsymbol{LP}}$	0.089	-0.069	0.178**	-0.083	0.048
	(1.063)	(0.765)	(0.568)	(0.137)	(0.046)
$\hat{\boldsymbol{\theta}}_{\boldsymbol{DH}}$	0.181	-0.141	0.006	na	0.097
	(0.128)	(0.011)	(0.055)		(0.068)
$\hat{\boldsymbol{\lambda}}$	-0.470***	-0.613**	-1.239	-0.747***	
CD	114.07**	86.23**	42.42***	42.41***	0.42*
$\vec{\Delta}$	11.347**				
Refer to table 3a for notes					

Turning to the long-run effects of RW, all but the CS-ARDL (2,2) estimates suggest a positive relationship between the female–male wage ratio and aggregate consumption for our entire panel. However, only the one-lag CS-ARDL specification returns a positive and significant estimate for RW under the extended regression. More interestingly, we notice that when two lags are imposed on the CS-ARDL approach, a negative coefficient is observed, even though the estimates are not significant. While one cannot draw firm conclusions on this phenomenon (given that the estimates are not statistically significant in this case), it is not unexpected that the persisting effect of a closing of the gender wage gap may negatively affect consumption, due to an increasing effect on saving after an initial period (Seguino and Floro 2003).

5.1.2. Short-run Dynamics

Table 4 reports the short-run results of the MG estimations for the baseline and extended regressions, respectively. The baseline CS-ARDL (1,1) short-run dynamics indicate that income and wage share have a positive and significant effect on aggregate consumption. As expected, the RW coefficient is positive but not significant in the CS-ARDL case.

Table 4: Short-run Regression Results for the Consumption Function (equation [2])

	ARDL		CS-ARDL	
Lags	(1,1)	(2,2)	(1,1)	(2,2)
Regressions with key variables				
$\hat{x}_Y$	0.148*** (0.042)	0.324*** (0.054)	0.039* (0.1.06)	0.315*** (0.067)
$\hat{x}_{RW}$	0.454 (0.385)	0.134 (0.305)	0.348 (0.278)	0.196 (0.158)
$\hat{x}_{WS}$	0.113 (0.204)	0.077 (0.229)	0.098* (0.207)	-0.147** (0.059)

Investment: In this subsection, a logarithmic formulation of the investment function is estimated, following closely the approach adopted in Stockhammer, Onaran, and Ederer (2008). In that work, the estimated investment equation is derived as an approximate specification of the general Kaleckian double-sided relation between investment and profits (the inverse of the wage share) around a dynamic equilibrium.

This general investment function is specified as:

$$lnI_{it} = f(\, lnY_{it}, lnRW_{it}, lnWS_{it}, lnINT_{it}) \qquad (12)$$

where gross investment, I_{it}, is the sum of residential and private investments, excluding government investment. Tables 5 and 6 report the long- and short-run coefficient estimates of the explanatory variables in the above investment function. Long-run estimates from the investment function are summarized in tables 5a and 5b.

5.2.1. Estimates of Long-run Effects

According to table 6a, the adjustment coefficients for all three panels ($for\ \rho = 1,2$) have the correct negative sign (except for the CS-ARDL [2,2] result) and are statistically significant, which implies that cointegration exists between the variables. The estimated coefficients for income, female-to-male wage ratio, and the wage share have the expected signs across all the different specifications. The size of the long-run estimates of income (Y) corresponds well with the existing literature and the theoretical underpinnings regarding the relationship between aggregate income and investment.

The estimates available from the different specifications also suggest the presence of a negative long-run distributional effect, as significant estimates of the average degree of responsiveness of the wage share for the entire panel of countries ranges from -0.19 to -0.21 when $\rho = 1$. Our CS-ARDL (1,1) estimate (-0.19) does however seem to be much lower than previous estimates available in Stockhammer, Onaran, and Ederer (2008) and Onaran and Galanis (2012). It should be noted that the estimated numerical coefficient of WS depends on a number of factors that may not have been accounted for in previous estimations—one of which may be the degree of gender equality.

While the main intuition of our framework suggests that an increase in the degree of gender wage equality may have an ambiguous effect on investment,[11] we notice that all ARDL and DL specifications point to a negative effect of RW on investment in the long run in our baseline estimation. However, the interest rate coefficient under the ARDL and DL specifications does produce ambiguous results in terms of sign; no significant results are found for the interest rate variable.

Table 5a: Baseline Regression Results for the Investment Function (equation [3])

	ARDL		**CS-ARDL**		**CS-DL**
Lags	(1,1)	(2,2)	(1,1)	(2,2)	ρ=1
	Regressions with key variables				
$\hat{x}_Y$	1.293*** (0.152)	2.918*** (0.828)	1.059*** (0.018)	1.318*** (0.027)	2.605*** (0.350)
$\hat{x}_{RW}$	-0.388 (2.102)	-3.190 (1.560)	-0.607*** (0.126)	-0.117 (0.096)	-0.078 (0.894)
$\hat{x}_{WS}$	-0.879 (0.630)	-0.704** (1.352)	-0.194** (0.128)	-0.037 (0.096)	-0.214* (0.271)
$\hat{x}_{INT}$	-0.017 (0.052)	0.581 (0.397)	-0.085 (0.039)	0.0004 (0.013)	-0.054 (0.055)
$\hat{\lambda}$	-0.591***	-0.747***	-0.079***	0.006	
CD	101.48	92.49***	9.78***	4.80***	-0.23
$\vec{\Delta}$	23.787**				

Refer to table 3a for notes

[11] See section 1.4.2.1

Robustness to additional explanatory variables: The baseline investment function is extended by including variables previously highlighted as potential determinants of the level of investment. The estimated long-run coefficient of aggregate income remains positive under the ARDL models and the CS-ARDL (1,1) model. However, parameter estimates for the other explanatory variables show very little sensitivity to the baseline results with respect to their benchmark assumptions. We do however find a consistently negative effect of RW and WS on the level of investment when lag $= 2$.

Although the estimated adjustment coefficients are qualitatively and quantitatively similar, the results in table 5b do not provide any consistent evidence of the effect of changes in household and business debt on the level of residential and private investment. Another important result is that, across all specifications, the CD test statistics are large enough that the null hypothesis of cross-sectional independence is strongly rejected at the 1 percent level in all cases. CS-DL results are omitted from the extended investment regression due to the increased number of regressors, which substantially reduces the degrees of freedom even with the suggested truncated lag length of one.

Table 5b: Extended Regression Results for the Investment Function (equation [3])

	ARDL		CS-ARDL			
Lags	(1,1)	(2,2)	(1,1)	(1,1)	(2,2)	(2,2)
Regressions with key variables						
$\hat{\theta}_Y$	1.282***	1.724***	0.709	1.209	0.941***	-0.269
	(0.326)	(0.233)	(0.795)	(2.381)	(0.011)	(0.175)
$\hat{\theta}_{RW}$	-1.586	-0.183	3.532	-0.279	-0.093**	-0.287***
	(3.435)	(0.347)	(3.389)	(0.436)	(0.047)	(0.092)
$\hat{\theta}_{WS}$	1.926	0.053	1.394	0.039	-0.253***	-0.089**
	(1.983)	(0.018)	(1.333)	(0.096)	(0.065)	(0.043)
$\hat{\theta}_{INT}$	0.029	0.046	0.005	0.041	-0.046***	2.696***
	(0.058)	(0.455)	(0.157)	(0.042)	(0.013)	(0.307)
$\hat{\theta}_{DH}$	-0.248	0.067***	1.404	1.209	0.328***	
	(0.257)	(0.005)	(1.161)	(2.381)	(0.025)	
$\hat{\theta}_{DB}$	-0.121	-0.009		0.045**		0.219***
	(0.209)	(0.534)		(0.023)		(0.030)
$\hat{\lambda}$	-0.795***	-0.188**	-0.647***	-0.084***	-0.178**	-0.046***
CD	44.62***	81.73***	78.61***	54.55***	16.86***	80.29***
Refer to table 3a for notes						

5.2.2 Short-run Dynamics

The estimates of the short-run country-specific ECMs provide evidence suggesting that long-run and short-run income effects on investment are mostly similar in the ARDL and CS-ARDL specifications.

In table 6, using the ARDL approach, we find that a higher level of gender wage equality has a positive (but not statistically significant) effect on investment in the short-run; the opposite is true of the CS-ARDL specifications, where negative estimates are observed, as with our long-run estimates in table 5, above. We also find that across the baseline results in table 6, the estimated short-run coefficients of the labor share of income are largely inconsistent in terms of the nature and direction of their association with aggregate investment. However, only negative coefficients on WS appear to be statistically significant in the short run, as observed using the CS-ARDL (2,2). These results for WS are consistent with the initial assumption of the potentially ambiguous effect of WS on investment. We also find no response of aggregate investment to short-run changes in interest rates, as in our long-run case. This result for the interest rate does seem to be in line with Kalecki's considerations that the interest rate is a less significant factor than the effect of aggregate profitability on the level of investment.

Table 6: Regression Results for the Investment Function (equation [3])

	ARDL		CS-ARDL	
Lags	(1,1)	(2,2)	(1,1)	(2,2)
Regressions with key variables				
$\hat{x}_Y$	1.417*** (0.227)	2.303* (0.045)	0.635 (0.757)	1.266* (0.734)
$\hat{x}_{RW}$	1.171 (0.909)	0.848 (1.354)	-9.323 (7.428)	-0.742 (0.467)
$\hat{x}_{WS}$	0.429 (0.563)	-0.032 (0.413)	-0.647 (0.802)	-1.574** (0.790)
$\hat{x}_{INT}$	0.017 (0.024)	0.017 (0.023)	0.045 (0.065)	-0.225 (0.188)

5.3. Foreign Sector

To model the foreign sector, separate import and export equations are estimated, as in Stockhammer and Wildauer (2016). In so doing, we are able to model imports as a function of domestic income (Y) and exports (exports are also taken to be a function of foreign income, FY). For our estimations, we employ the following export (X) and import (M) functions:

$$lnX_{it} = f(\, lnRW_{it}, lnWS_{it}, lnEX_{it}, lnFY_{it}) \tag{13}$$

$$lnM_{it} = f(lnY_{it,}lnRW_{it}, lnWS_{it}, lnEX_{it}, lnX_{it}) \tag{14}$$

5.3.1 Exports

Estimates of long-run effects: The results of the export function regression in table 7 indicate that the error correction coefficients fall within the dynamically stable range (being statistically significant and negative), and therefore the null hypothesis of no long-run relation is rejected. This finding indicates that there is compelling evidence for conditional convergence to country-specific dynamic equilibriums in our sample of 46 countries.

Table 7: Regression Results for the Export Function (equation [4])

	ARDL		**CS-ARDL**		**CS-DL**
Lags	(1,1)	(2,2)	(1,1)	(2,2)	ρ =1
Regressions with key variables					
$\hat{x}_{RW}$	1.603*** (0.444)	1.486*** (0.433)	1.278*** (0.290)	0.689** (0.331)	1.519*** (0.006)
$\hat{x}_{WS}$	-0.660*** (0.171)	-0.553*** (0.159)	-1.072*** (0.238)	-1.648*** (0.290)	-1.465** (0.407)
$\hat{x}_{EX}$	-0.217*** (0.068)	-0.193*** (0.063)	-0.507*** (0.076)	-0.131* (0.071)	-0.538** (0.228)
$\hat{\lambda}$	-0.084***	-0.046***	-0.235***	-0.315***	
CD	84.00***	92.49***	12.44***	8.67**	2.02**
$\vec{\Delta}$	9.008*				
Refer to table 3a for notes					

Furthermore, results from table 7 suggest that in the long run, the wage share (WS) and exchange rates (EX) are negatively associated with the level of exports, as expected. However, significant positive effects of the female-to-male wage ratio are observed on total exports for the entire panel. Also, there is evidence of cross-sectional dependence under all specifications—though the CD statistics appear more marginal for the CS estimates. It is also worth noting that the reported results in table 7 are consistent across the different model specifications. More importantly, these findings for wage share and exchange rates follow our a priori expectations and are also in line with the results of Stockhammer and Wildauer (2016) and Onaran and Obst (2016). The results for RW suggest that increasing women's wages may positively affect the level of exports; we do however expect that this result may be different in the short run.

Short-run dynamics: The short-run dynamics for the export function are reported in table 8. We find no significant short-run association between the female–male wage ratio and the level of exports. We ascribe the export's lack of a short-run reaction to changes in RW primarily to the preeminence of developed countries in our panel sample, as we expect that the level of exports in developing countries responds more swiftly to changes in the level of gender wage equality in comparison to more developed economies.

Table 8: Regression Results for the Export Function (equation [4])

	ARDL		CS-ARDL	
Lags	(1,1)	(2,2)	(1,1)	(2,2)
Regressions with key variables				
$\hat{x}_{RW}$	0.036	0.050	-0.469	-0.462
	(0.095)	(0.108)	(0.808)	(0.775)
$\hat{x}_{WS}$	-0.239***	-0.226**	-0.449*	-0.075
	(0.074)	(0.075)	(0.5110	(0.469)
$\hat{x}_{EX}$	-0.082**	-0.083**	-0.054	-0.250**
	(0.037)	(0.037)	(0.065)	(0.093)

5.3.2 Imports

Estimates of long-run effects: Table 9 presents the estimated results of the import equation for the entire panel. In summary, the impact of domestic income on the level of imports is positive and statistically significant in most cases in the baseline and extended regressions.

However, the results for the effect of the wage share and exchange rates are ambiguous, as statistically significant effects are observed in opposite directions in both tables. For the female-to-male wage ratio, we only find our estimated coefficients to be significant when pointing to a positive relationship between RW and imports; the theoretical literature on this issue is also by and large inconclusive but clearer inference is expected from the results for the different country groups.

Table 9: Regression Results for the Imports Function (equation [4])

	ARDL		CS-ARDL		CS-DL
Lags	(1,1)	(2,2)	(1,1)	(2,2)	ρ =1
Regressions with key variables					
$\hat{x}_Y$	2.094***	1.543***	1.638	1.493***	1.608***
	(0.277)	(0.035)	(1.450)	(0.132)	(0.296)
$\hat{x}_{RW}$	-2.252	1.201***	0.461***	3.722	-1.165
	(1.444)	(0.105)	(0.144)	(1.045)	(1.835)
$\hat{x}_{WS}$	1.109**	0.628***	-1.104	-0.049	-1.214
	(0.564)	(0.067)	(1.864)	(0.121)	(2.221)
$\hat{x}_{EX}$	0.167	0.107**	-1.186	-0.040	-0.062**
	(0.208)	(0.048)	(1.174)	(0.049)	(0.138)
$\hat{\lambda}$	-0.595***	-0.216***	-0.709***	-0.031**	
CD	67.22	84.33***	12.45***	19.12***	1.35**
$\vec{\Delta}$	98.766***				
Refer to table 3a for notes					

Short-run dynamics: Short-run findings for our import function are closely similar in magnitude and direction to our long-run results. For the sake of brevity, we find a positive effect of domestic income and exchange rates on imports, following our benchmark assumptions. We also find, as in our long-run case, that the wage share and ratio of female–male wages have no clearly identified direction of association with the level of imports for all 46 countries in our panel.

Table 10: Regression Results for the Imports Function (equation [4])

	ARDL		CS-ARDL	
Lags	(1,1)	(2,2)	(1,1)	(2,2)
Regressions with key variables				
$\hat{x}_Y$	0.803***	1.506***	0.814*	1.405***
	(0.170)	(0.165)	(0.439)	(0.152)
$\hat{x}_{RW}$	1.614	-0.882	0.330**	-1.987
	(1.213)	(1.228)	(0.129)	(2.658)
$\hat{x}_{WS}$	0.022	0.953	-0.866	-0.015
	(0.178)	(0.801)	(0.999)	(0.104)
$\hat{x}_{EX}$	0.106	0.109	-0.561	-0.026
	(0.099)	(0.100)	(0.656)	(0.039)

5.4. Total Effects: Demand-led Regimes

Using the annual panel dataset, the preferred CS-ARDL approach was employed to account for endogeneity, cross-country heterogeneity, and cross-sectional dependence that arise from unobserved common factors. The main findings suggest a general long-run cointegrating relationship between the relevant explanatory and dependent variables of the various equations of the AD components. Furthermore, these results from the formal statistical analyses enable one to infer the long-run and short-run elasticities of the regressors in each of the equations for the macroeconomic and net export equilibrium conditions.

To examine the distributional effects on AD as well the gendered impact of income inequality on the macro economy, one needs to determine, from the estimated elasticities, the partial effects of the wage share on the various components of AD. These converted marginal effects can then be used to account for the cumulative effects of the wage share on AD.[12] Table 11 reports the long-run and short-run marginal effects for the entire panel. Marginal effects and cumulative effects are also reported for different country groups using the conversions from the CS-ARDL (1,1) model.

Table 11 reports the marginal effects of the wage share on the components of AD and its cumulative effect on AD based on average parameters. Our results indicate overall long-run wage-led growth and short-run profit-led growth following the respective signs of private excess demand (Y^{PED}).[13]

Table 11: **Marginal Effects of a 1 Percent Change in WS on Private Excess Demand for the 46-country Panel**

X	**1 lag; 3 CS lags**		**2 lags; 2 CS lags**	
	LR	**SR**	**LR**	**SR**
Consumption	0.671	0.133	1.129	-0.199
Investment	-0.049	-0.163	-0.009	-0.397
Net exports	-0.11	-0.016	-0.288	-0.014
$\mathbf{Y^{PED}}$	0.512	-0.046	0.832	-0.610
Multiplier	4.484	1.156	4.149	1.603
Total effect	2.296	-0.053	3.452	-0.978

Notes: Long run: wage-led growth || Short run: profit-led growth

This finding is in line with Sánchez and Luna's (2014) results for Mexico and may also corroborate Blecker's (2016) argument that the magnitude of the impact of the wage and profits share on the components of AD may depend substantially on the time period examined, which, in turn, may explain

[12] Elasticities are converted into marginal effects using:

$$\frac{\partial Y}{\partial WS} = \hat{\beta}_{C,WS}\left(\frac{C}{WS}\right) + \hat{\beta}_{I,WS}\left(\frac{I}{WS}\right) + \hat{\beta}_{X,WS}\left(\frac{X}{WS}\right) + \hat{\beta}_{M,WS}\left(\frac{M}{WS}\right)$$

[13] Private excess demand as defined by Onaran and Galanis (2012) is the sum of the partial effects of distribution of demand prior to the multiplier effects.

the conflicting results present in the literature on the impact of distribution on AD: "Some distributional effects may be more important in the short run (over a few quarters or years, or the length of an ordinary business cycle), while others are likely to be more important in the long run (across, say, one or more decades). In particular […], the positive effects of higher profit shares (or lower labor costs) on investment and net exports are mainly short-run phenomena, while the sensitivity of workers' consumption to their wage income is, if anything, likely to be stronger in the long run" (Blecker 2016, 3).

Furthermore, the long-run and short-run multipliers are calculated using equation (6), above. The results suggest that the long-run multipliers (accumulated effects) are substantially larger than the reported interim (short-run) multipliers in the one- and two-lag periods. We note that the short-run multipliers are consistent with what is usually observed in the literature (Onaran and Galanis 2011; Keifer and Rada 2014).

With regard to the long-run multipliers, we do not have sufficient research to compare our findings, but it is important to note that the relevant long-run coefficients (elasticities) appear statistically significant more frequently than the short-run results do.[14] Finally, the last row of table 11 shows the total effect on AD from a change in income distribution when the multiplier mechanism is accounted for. As expected, there is a larger impact on equilibrium income from changes in the wage and profit share when the multiplier is introduced.

5.4.1 Subsample Cumulative Effects

Evidence from the panel results suggests a long- and short-run relationship between functional income distribution and AD, with demand-led growth observed to be profit-led in the short run and wage-led in the long run. However, this result for demand-led growth may vary for certain country groups; in what follows, the marginal and cumulative effects are reported for the different country groups. As stated in earlier sections, the 46 countries in the panel are grouped into high-income, middle-income/SIEs, and low-income/LIAEs countries. Marginal effects are calculated using the same formula as in the panel

[14] Both statistically significant and nonsignificant point estimates are employed in the calculation of the partial and cumulative effects due to the scant evidence of statistical significance of the short-run estimates. However, the lack of statistical significance of some of these estimates implies a lack of precision in the estimation of the partial effects and as such, we are not entirely certain that the value of the corresponding parameter in the underlying regression model is zero.

case; however, the elasticities employed in the calculation are derived from taking the average of the country-specific MG estimates for the countries in each income group. Similarly, in this case, the consumption share is also derived from the subsample weighted sum; these subsample averages are calculated using the GDP-weighted averages of the various AD components (C, I, X, M) normalized by income and the GDP-weighted average of the wage share.[15]

Table 12: Long-run WS Effects on AD for Different Income Groups

	Panel	Developed countries	SIEs	LIAEs
Consumption	0.671	-0.661	2.879	4.307
Investment	-0.049	3.701	-0.274	0.169
Net export	-0.11	-0.493	0.498	-0.154
$\mathbf{Y}^{PED}$	0.512	2.547	3.103	4.322
Multiplier	4.484	0.561	1.769	0.563
Total effect	2.296	1.429	1.489	2.433
Openness	7 percent	21 percent	36 percent	61 percent

Tables 12 and 13 report the results, where only the partial, cumulative, and multiplier effects are presented for the long- and short-run cases using the estimates from the preferred CS-ARDL (1,1) model. Notably, the error correction coefficient is negative and significant across all regressions, indicating cointegration between the regressors and their respective AD components.

Several studies point to the importance of economic structure and openness to trade in determining the nexus between income distribution and aggregate income (see, for example, Blecker [2016] and Stockhammer Onaran, and Ederer [2008]). In the same vein, Onaran and Galanis (2012) show that economies with a higher degree of openness are more likely to be profit-led than wage-led—at least in the short run. This is perhaps because net export effects can be directly associated with the openness index (i.e., on the share of exports and imports in GDP), such that small, open economies, for example, may be more sensitive to net export volatility, thus potentially leading to circumstances where the negative effects of a rise in the wage share are enough to overshadow any positive wage share effects

[15] The cumulative effect for each country group is derived thus:

$$\frac{\partial \boldsymbol{Y}}{\partial \boldsymbol{WS}} = \widehat{\boldsymbol{\beta}}_{C,WS}\left(\emptyset\frac{\boldsymbol{C}}{\boldsymbol{Y}}\right)\frac{\mathbf{1}}{\emptyset \boldsymbol{WS}} + \widehat{\boldsymbol{\beta}}_{I,WS}\left(\emptyset\frac{\boldsymbol{I}}{\boldsymbol{Y}}\right)\frac{\mathbf{1}}{\emptyset \boldsymbol{WS}} + \widehat{\boldsymbol{\beta}}_{X,WS}\left(\emptyset\frac{\boldsymbol{X}}{\boldsymbol{Y}}\right)\frac{\mathbf{1}}{\emptyset \boldsymbol{WS}} + \widehat{\boldsymbol{\beta}}_{M,WS}\left(\emptyset\frac{\boldsymbol{M}}{\boldsymbol{Y}}\right)\frac{\mathbf{1}}{\emptyset \boldsymbol{WS}}$$

As in the previous marginal effects conversion, the elasticities are converted to marginal effects and then normalized by income. However, for the various country groups, these elasticities are calculated as GDP-weighted averages for the countries in each subpanel.

on investment and consumption—this is found to be the case in the short run for the low-income countries (see table 13).

The results in table 12 show the long-run effects for the different country groups. The results suggest that growth is wage-led for all country groups and the entire panel. However, there are several interesting patterns. First, the evidence seems to suggest a negative wage share effect on net exports for the high- and low-income countries. The results for the high-income countries follow the postulation that large and relatively closed economies (which are more likely to be high-income countries) have a small net export effect compared to their consumption effect and, as such, are strongly wage-led. In LIAEs, the evidence also points a small net export effect. However, for LIAEs, the observed net export effect is relatively smaller compared to the other country groups and, more importantly, this effect is always smaller than the consumption effect—hence the observed long-run wage-led growth regime. Secondly, following from the first observation, the degree of openness may not always be directly linked with the demand regime of an economy when the level of development and the long-run time dimension comes into play. Third, the wage share effect on consumption for high-income countries is surprisingly negative overall and becomes larger and positive with lower levels of aggregate income (i.e., in SIEs and LIAEs).

There is also a considerably large positive effect on investment from changes in the wage share for the high-income countries in our panel. This large positive effect is what determines the wage-led regime for this country group and not the wage effect on consumption, as would be expected; in fact, there is a negative consumption effect for this group of countries. This relationship may be explained by the argument that, in a demand-constrained society, potentially rising wages can serve as a boost to the long-term profitably of firms by stimulating investment (Schelling 1946; Bhaskar 1992; Onaran and Yenturk 2001) However, there is no comprehensive answer to the question of why corporate investment and profitability will be positively affected by a higher wage share, but the positive impact of increasing wages on residential investment is a possible avenue.

There is a negative wage share effect on investment for the SIEs in our panel, suggesting that an increase in the wage share may indicate a loss of profit to capitalists or businesses—which may in turn discourage future investment.

Table 13: Short-run Effects on AD for Different Income Groups

	Panel	Developed countries	SIEs	LIAEs
Consumption	0.133	0.076	0.131	1.333
Investment	-0.163	0.453	-0.548	2.123
Net Export	-0.016	-0.119	-0.348	-4.004
Y^{PED}	**-0.046**	**0.410**	**-0.765**	**-0.548**
Multiplier	**1.156**	**1.675**	**0.718**	**1.783**
Total effect	-0.053	0.687	-0.549	-0.977

Our short-run results—reported in table 13—appear notably different from their long-run counterparts. Here, we find evidence to support a profit-led demand regime for the entire panel. This profit-led growth hypothesis is also confirmed for the SIEs and LIAEs. For the SIEs, there are negative WS effects on investment and net exports that overshadow the positive WS effect on consumption—hence a profit-led demand growth regime. We only find the coexistence of a long- and short-run wage-led growth regime for the high-income countries' group.

Table 13b: Growth Regimes Summarized

	Panel	DC	SICS	LIAES
Long run	Wage-led	Wage-led	Wage-led	Wage-led
Short run	Profit-led	Wage-led	Profit-led	Profit-led

Moreover, these findings follow the hypothesis that the economy's degree of openness directly corresponds to the net export effects. Following the benchmark assumptions, findings suggest that relatively closed economies (developed countries) are more strongly wage-led, while medium-sized open economies (SIEs) tend to have relatively smaller net export effects, leading them to often follow a profit-led growth regime. However, low-income economies—often small, open economies (as in our panel)—tend to have large negative net export effects (-4.004), which may result in the total demand regime becoming profit-led. These a priori expectations of a corresponding relationship between openness and the external sector seem to hold only for the short-run demand-led growth determination. Furthermore, the short-run multipliers, as earlier mentioned, are more consistent with the existing literature—these are especially similar to the multipliers derived in Onaran and Galanis (2012).

5.4.2 Gender Effects on the Macroeconomy

Having identified the growth regimes of the different country groups above, we now estimate the relative contributions of gender inequality to each of the macroeconomic aggregates. In doing so, the role of economic structure in affecting the relationship between gender equality and macroeconomic outcomes is accounted for. An obvious way to employ gender as an explanatory macroeconomic tool is to disaggregate at least one of the components of AD by gender. As previously described, such a disaggregation fits well with the Bhaduri and Marglin (1990) demand-led model, which emphasizes the distributional differentiation of economic agents.

To determine the macroeconomic effects of gender inequality, we investigate under what conditions the different country groups—and the entire panel—display a gender equality–led growth regime in the long or short run. As earlier explained, a gender equality–led growth regime is one that is gender cooperative and an increase in gender equality results in injections exceeding leakages (S +M < I +X) (i.e., a redistribution stimulates AD, leading to an increase in output). Tables 14 and 15 report the long- and short-run *RW* elasticities[16] for each of the individual AD components using the preferred CS-ARDL (1,1) model. At the panel level, the gender effects are quite modest and expansionary in the long run, but larger and contractionary in the short run. These magnitudes of effects are as expected, as gender effects are more likely to strongly persist in the short run than in the long run.

Table 14: Summary of Long-run Gender Effects

L.R	Panel	DC	SICS	LIAES
S	0.008	10.23	-0.377	6.743
M	0.461	-1.875	0.632	0.362
SUM	**0.469**	**8.355**	**0.255**	**7.105**
I	-0.607	1.826	-0.388	1.723
X	1.278	1.884	4.612	2.387
SUM	**0.671**	**3.71**	**4.224**	**4.11**
Effect	**E**	**C**	**E**	**C**

Note: where **E → expansionary effect** (gender cooperative) and **C → contractionary effect** (gender conflictive)

Table 14 presents the long-run directional effects of greater gender equality on the various components of AD. While the results for the entire panel suggests that greater gender equality exerts expansionary pressures on economic growth, this inference does not hold firmly when viewed through structural lenses. Particularly, greater gender equality imposes contractionary pressures on growth for developed and low-income countries, while it produces expansionary pressure on SIEs/middle-income countries.

Table 15 provides a summary of how lower gender wage gaps affect various macroeconomic outcomes in the short run and the implications for growth. As might be expected, in the short run, greater gender

[16] The calculated elasticities are employed whether or not they appear statistically significant in the demand-led regime determination above.

inequality is contractionary for the entire panel due to the potential negative effects of higher female wages on competitiveness. As in the long-run case, we observe varying results for the different country groups. In particular, the short-run impact of greater gender equality on the process of economic growth is contradictory for the developed and semi-industrialized countries but expansionary for the low-income countries.

Table 15: **Summary of Short-run Gender Effects**

S.R	PANEL	DC	SICS	LIAES
S	0.652	-1.446	-0.076	0.664
M	0.330	-0.317	-0.076	0.435
SUM	**0.982**	**-1.129**	**-0.152**	**1.099**
I	-9.323	0.534	-0.236	-0.636
X	-0.469	-0.664	-0.580	3.576
SUM	**-9.792**	**-0.13**	**-0.816**	**2.94**
Effect	**C**	**E**	**C**	**E**

A few things stand out about these results. First, we observe that the magnitude of the gender effect on savings and imports is much larger for LIAEs and developed economies both in the long and short run. However, while these gender effects are large enough to compel a contractionary squeeze on growth in developed economies in the long and short run, in LIAEs they are only large enough to drive contractionary pressures on growth in the long run.

Secondly, the results further underscore the crucial role played by gender equality in the relationship between equality and macroeconomic outcomes. As observed, gender wage equality appears to have differential effects in the long and short run depending on the nature of economic development and economic structure. For example, we find that in the short run, gender-based wage differentials are only linked to export growth in high- and low-income countries; however, the process by which this occurs in both country groups is entirely different. In the high-income group, we observe that this expansionary stimulus to short-run growth is mostly a result of the positive impact of greater gender wage equality on investment. However, the exact opposite is true for low-income countries in the short run, as the potential positive push on growth is due to the positive impact of greater gender equality on consumption, imports, and exports, with higher female wages negatively impacting investment.

Furthermore, in the long run, the results suggest that greater gender wage equality is contractionary for growth in both high-and low-income countries, implying direct opposite AD effects in the long and short run for low- and high-income countries. While these long-run effects are not unexpected for the low-income countries as they advance toward higher development, we are unsure as to the reasoning behind the potential contractionary pressures on growth due to sustained increases in female wages in high-income countries. However, Seguino and Braunstein (2019) highlight the concurrent decline in male labor force participation as women's increases and the impact of the quality of new jobs generated from growth on women's (and men's) position in the labor market. To some extent, this may be useful in explaining this short-run contractionary pressure on growth for LIAEs if improvements in female wages are combined with an overall fall in the wage share due to the impact on male workers or access to better jobs.

Table 16: **Summary of Gender Effects**

	Panel	**DC**	**SICS**	**LIAES**
Long-run effect	Gender cooperative	Gender conflictive	Gender cooperative	Gender conflictive
Short-run effect	Gender conflictive	Gender cooperative	Gender conflictive	Gender cooperative

In summary, the short-run results clearly follow a priori expectations. As stated in earlier sections, we expect that due to the adoption of export orientation in many middle-income or semi-industrialized economies, low female wages may indeed act as a stimulus to growth. A large body of literature has linked gender employment and wage discrimination embedded in social and labor market practices to increased export competitiveness, which is largely dependent on low unit labor costs, which women can readily provide under these conditions (Deyo 1989; Hsiung 1996; Seguino 1997; Blecker and Seguino 2002).

In addition, it is expected that greater gender wage equality will improve demand-led growth conditions in LIAEs and developed countries. In LIAEs, it is assumed that women are more largely concentrated in subsistence farming and domestic unpaid care work. As such, any improvement in female wages will lead to increased productivity in the agricultural and care sectors, as well in female labor force participation, all of which work to increase the productivity and human capital endowments of women and children (Darity 1995; Seguino 2012).

In contrast, the long-run results do not necessarily conform to any previous expectations. The most puzzling aspect of the long-run results is that they are exactly opposite of the observed short-run results for all the country groups and for the entire panel. While growth can be assumed to respond to expansionary pressures in the long run for the entire panel, only SIEs can be assumed to follow a similar pattern. However, contrary to expectations, the results suggest that this expansionary pressure on growth is mostly mobilized by the positive effect a sustained increase in female wages has on exports in the long run. This striking result may imply that while gender inequality is linked to export-led growth at a country's adoption stage of export orientation, a continual closing of the gender gap may have the long-term result of increasing export potentials due to the implications on labor productivity transmitted through rising female educational attainment, the diversification of the female labor force, and increased female bargaining power (see Seguino 2013). This result may buttress the neoclassical theory that posits that at long-run full employment,[17] women's wages are expected to rise to indicate their increased productivity. However, even when such rises occur, there is still evidence of an unequal distribution of income among substitute workers of different genders.

6. CONCLUSION

A revived interest in the relationship between income distribution and the macroeconomy is evident in the academic literature and amongst policymakers (Stockhammer 2011, 2013b; Onaran and Galanis 2013). The sustained deceleration in global economic growth and the global financial crisis of 2007–9 has also renewed interest in this strand of economic research, with emphasis on debunking the assumed economy-wide benefits of mainstream policy recommendations that favor wage moderation and austerity (Rajan 2010; Wisman and Baker 2011; Stiglitz 2010; Wade 2011a; Hein, Truger, and van Treeck 2012; Wisman 2013b).

Insight provided from this large body of research also suggests that gender inequality may have substantial implications for economic growth. However, not much of the existing literature has focused on the channels through which gender inequality affects economic growth. Consequently, research

[17] Recall that in the typical usage of the "long run" in this study, we refer to the quasi-equilibrium that implies a succession of many short-run periods. Also, in reference to growth effects, we are concerned with the potential pressures on—or against—economic growth, as is implied in the Kaleckian demand-led growth framework.

originating from revitalized academic concentration on the determinants of AD and output are essentially bereft of a gendered perspective, which, even when present, is rarely empirically estimated. This study aims to close this gap in the literature by investigating the effect of income distribution and gender equality on each of the variables in the macroeconomic and net export equilibrium conditions.

Preliminary findings suggest that growth is profit-led in the short run and wage-led in the long run. However, we find wage-led growth in the long and short run for high-income countries. In addition, following the estimations of gender effects, we are able to posit that global economic growth is gender equality–led and wage-led in the long run; this is also the case for middle-income countries. Furthermore, short-run improvements in gender wage equality are consistent with profit-led growth in low-income countries, suggesting that greater gender equality makes the growth regime for LIAEs more wage-led or less profit-led in the long run.

Furthermore, the results highlight the role of distribution in determining (private domestic) demand and the effect this may have on employment and growth. For example, given that global growth is wage-led, even for countries at different stages of economic development, one can assume that wage moderation is unlikely to stimulate employment in the long run.

The findings also suggest that a redistribution in favor of wages, away from profits, may create a path to sustainable growth due its increased effect on demand and technological progress; such a path to sustainable growth will also need a rebalancing in the form of gender equality which—as evident in the literature—promotes productivity and long-run growth, and drives a wage-led growth agenda. Such a wage-led growth strategy that aims at an economic regime that highlights increased demand and innovation as crucial goals will require strong policy coordination. Further research is needed to fully understand why increasing women’s wages elicits different responses from macroeconomic aggregates in the long and short run.

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APPENDIX

Table A1: Country Classification and Abbreviations (World Bank)

High-income countries		**Middle-income countries**		**Low-income countries**	
Country	*Abbr.*	*Country*	*Abbr.*	*Country*	*Abbr.*
Australia	AUS	**Argentina**	ARG	**Belize**	BLZ
Austria	AUT	**Botswana**	BWA	**Burkina Faso**	BFA
Belgium	BEL	**Brazil**	BRA	**Egypt**	EGY
Canada	CAN	**Bulgaria**	BGR	**Gambia**	GMB
Chile	CHL	**Colombia**	COL	**Philippines**	PHL
Denmark	DNK	**India**	IND	**Senegal**	SEN
Finland	FIN	**Macedonia**	MKD	**Sri Lanka**	LKA
France	FRA	**Mauritius**	MAU	**Thailand**	THA
Germany	DEU	**Mexico**	MEX	**Tunisia**	TUN
Greece	GRC	**Peru**	PER	**Uganda**	UGA
Ireland	IRL	**Poland**	POL		
Italy	ITA	**Turkey**	TUK		
Japan	JPN				
Korea	KOR				
Macao, China	MAC				
Netherlands	NLD				
New Zealand	NZD				
Norway	NOR				
Singapore	SGP				
Spain	ESP				
Sweden	SWE				
Switzerland	CHE				
United Kingdom	GBR				
United States	USA				

Table A2: Data Definitions and Sources

Abbreviation	Full Variable Name	Unit	Source
Y	GDP at 2010 market prices (NY.GDP.MKTP.KD)	billion, US dollars 2010 = 1	WDI
C	Final consumption expenditure at 2010 market prices (NE.CON.TOTL.KD)	billion, US dollars 2010 = 1	WDI
I	Gross fixed capital formation at 2010 market prices (NE.GDI.FTOT.KD)	billion, US dollars 2010 = 1	WDI
X	Exports of goods and services at 2010 prices (NE.EXP.GNFS.KD)	billion, US dollars 2010 = 1	WDI
M	Imports of goods and services at 2010 prices (NE.IMP.GNFS.KD)	billion, US dollars 2010 = 1	WDI
WS	Adjusted wage share: total compensation, as a percentage of gross value-added at factor cost (ALCD2/NY.GDP.FCST.CN/LFS/IMF)	percent GDP	IMF
RW*	Ratio of female to male median weekly wages (percent)	(percent national estimate)	ILO (LABORSTA)
INT	Real long-term interest rates, GDP deflator	percent per annum	IMF
EX	Nominal effective exchange rate	2010 = 1	BIS and FRED
LP	Female labor force participation rate (SL.EMP.WORK.FE.ZS)	percent	WDI
DH	Household debt-to-GDP ratio, adjusted for breaks	percent	BIS
DB	Business debt-to-GDP ratio, adjusted for breaks	percent	BIS
FY	Foreign income (global GDP)	billion, US dollars 2010 = 1	WDI

Note: For this study, RW data for all countries (with the exception of Burkina Faso, Sri Lanka, and Senegal) is collected from the old ILO databank, LABORSTA. This has recently been integrated with the new database, ILOSTAT. For Burkina Faso and Senegal, we use the "wage equality for similar work" survey data collected and reported yearly by the World Economic Forum; for Sri Lanka, we use ILO data on male and female average nominal wages to calculate the RW ratio.

Table A3: Summary Statistics

Variable		Mean	Std. Dev.	Min	Max	Observations
year	overall	2000	8.94741	1985	2015	N = 1426
	between		0	2000	2000	n = 46
	within		8.94741	1985	2015	T = 31
Y	overall	8.86e+11	2.02e+12	3.17e+08	1.67e+13	N = 1412
	between		1.98e+12	6.45e+08	1.23e+13	n = 46
	within		4.69e+11	-3.68e+12	5.35e+12	T-bar = 30.6957
C	overall	6.93e+11	1.65e+12	3.27e+08	1.38e+13	N = 1411
	between		1.61e+12	6.33e+08	1.02e+13	n = 46
	within		3.83e+11	-3.06e+12	4.32e+12	T-bar = 30.6739
I	overall	1.95e+11	4.29e+11	8.37e+07	3.40e+12	N = 1411
	between		4.17e+11	1.71e+08	2.46e+12	n = 46
	within		1.13e+11	-7.54e+11	1.14e+12	T-bar = 30.6739
X	overall	1.88e+11	2.87e+11	1.08e+08	2.21e+12	N = 1409
	between		2.52e+11	2.20e+08	1.25e+12	n = 46
	within		1.40e+11	-6.64e+11	1.15e+12	T-bar = 30.6304
M	overall	1.90e+11	3.25e+11	2.33e+08	2.81e+12	N = 1409
	between		2.89e+11	3.60e+08	1.65e+12	n = 46
	within		1.53e+11	-8.65e+11	1.36e+12	T-bar = 30.6304
EX	overall	132.5066	315.0494	33.37	9392.84	N = 1263
	between		117.952	79.18355	832.2225	n = 45
	within		295.7287	-650.8259	8693.124	T-bar = 28.0667
INT	overall	30.64488	361.0631	-.14	9394.293	N = 1254
	between		133.1602	1.950699	910.2247	n = 46
	within		332.7064	-871.7727	8514.713	T-bar = 27.2609
LP	overall	50.85397	13.3549	18.709	82.333	N = 1222
	between		12.91041	21.36378	80.58048	n = 46
	within		3.762701	34.57611	62.72463	T-bar = 26.5652
RW	overall	66.69681	15.91788	22.43124	94.18137	N = 1247
	between		15.45325	25.592	90.36703	n = 46
	within		4.797519	47.21651	83.45574	T-bar = 27.1087
WS	overall	54.12562	11.14595	25.4638	90.5	N = 1287
	between		10.91147	33.64105	85.07213	n = 46
	within		4.528517	39.24498	88.61873	T-bar = 27.9783
DH	overall	45.65888	27.9053	-.0019099	139.5	N = 1247
	between		26.05966	4.926087	112.3833	n = 46
	within		13.63355	2.821843	91.56541	T-bar = 27.1087
DB	overall	61.4459	39.21051	2.013643	264.9	N = 1241
	between		37.92482	8.97934	168.7467	n = 46
	within		15.82774	-26.30076	157.5992	T-bar = 26.9783

FY	overall	5.11e+13	1.31e+13	3.16e+13	7.52e+13	N = 1426
	between		0	5.11e+13	5.11e+13	n = 46
	within		1.31e+13	3.16e+13	7.52e+13	T = 31
NX	overall	-.0861276	.5079779	-4.693875	.5933776	N = 1407
	between		.4796666	-2.481668	.2982423	n = 46
	within		.1842213	-2.744342	.9466142	T-bar = 30.587

Note: The low min values for LP and RW are from Egypt, Tunisia, and Turkey