J. Int. Financ. Markets Inst. Money 78 (2022) 101541

Contents lists available at ScienceDirect

Journal of International Financial Markets, Institutions & Money

journal homepage: www.elsevier.com/locate/intfin

# Dynamic relationship between exchange rates and stock prices for the G7 countries: A nonlinear ARDL approach

Salah A. Nusair [a], Dennis Olson [b,*]

[a] DeGroote School of Business–Finance Area, McMaster University, Mississauga, Ontario, Canada
[b] College of Business, Zayed University, Abu Dhabi, United Arab Emirates

ARTICLE INFO

*JEL classifications:*
F31
G11
G15

*Keywords:*
Exchange rates
Stock prices
Asymmetry
Nonlinear ARDL model

ABSTRACT

This paper employs linear and nonlinear ARDL models to examine the short-run and long-run relationship between stock prices and exchange rates in the G7 countries. Both the flow-oriented approach that exchange rates affect stock prices and the portfolio balance approach that stock prices affect exchange rates are supported in the short-run. Neither model is supported in the long-run using linear ARDL models, but the nonlinear ARDL model shows evidence supporting the portfolio balance approach in four of the countries. In these four countries we find that rising and falling stock prices have significant long-run effects on their exchange rates. Furthermore, Granger causality tests confirm that causality runs from stock prices to exchange rates in six of the countries. Thus, the use of a longer and more recent data set provides stronger long-run support for the portfolio balance approach than found in most of the recent literature, while we confirm results of recent research showing no long-run evidence of causation running from exchange rates to stock prices.

## 1. Introduction

The link between stock prices and exchange rates has long been the subject of research and policy debates and it has received considerable attention in the financial economics literature. A significant relationship between stock prices and exchange rates could suggest that shocks in one market are transmitted quickly to the other market. Such a relationship has important implications for economic policies and international portfolio and capital budgeting investment decisions (Chkili and Neguyen, 2014). For example, if exchange rates affect stock prices, controlling exchange rates might help to stabilize stock markets. Likewise, if stock prices affect exchange rates, then policies that stabilize stock prices may reduce exchange rate volatility. Also, if the two markets are connected, investors may be able to use information available in one market to predict the other market for hedging and speculation purposes.

The two major theories that explain the interactions between foreign exchange markets and stock markets are the flow-oriented (goods market approach) and the stock-oriented (portfolio balance approach).[1] The flow-oriented approach, as derived from the work of Dornbusch and Fisher (1980), states that exchange rates affect stock prices. For example, Aggarwal (1981) postulates that currency appreciation reduces international competitiveness, while depreciation enhances competitiveness—leading to a negative relationship between the two variables. However, Bahmani-Oskooee and Saha (2016) argue that the traditional international competitiveness story holds for export-oriented firms, but that exchange rate movements have the opposite impact on domestic import-

* Corresponding author.
*E-mail address:* dolson_prof@yahoo.com (D. Olson).

[1] The flow-oriented and the portfolio balance approaches were initially developed as models of exchange rate determination.

https://doi.org/10.1016/j.intfin.2022.101541
Received 3 June 2020; Accepted 4 March 2022
Available online 11 March 2022

oriented firms. Following a currency depreciation, stock prices increase (decrease) when a country has more (less) export-oriented firms than import-oriented firms, while for currency appreciations, stock prices increase (decrease) when a country has less (more) export-oriented firms than import-oriented firms. Thus, the flow-oriented approach posits that causality runs from exchange rates to stock prices and that the relationship that can be either positive or negative. In contrast, the stock-oriented models of exchange rates, such as Frankel (1983), suggest that causation runs from stock prices to exchange rates and posits a positive relationship between the two variables.[2] To illustrate, Bahmani-Oskooee and Saha (2018) argue that stock prices affect exchange rates through wealth and expectation effects channels. A rising domestic stock market increases investors' wealth, causing the demand for money to increase, which leads to domestic currency appreciation. Also, rising interest rates attract foreign capital inflows, which results in further stock price increases and currency appreciation.

Despite the large body of research examining the link between stock prices and exchange rates, the empirical evidence is quite mixed about the nature of the causality relationship. As a possible explanation, Bahmani-Oskooee and Sohrabian (1992) have suggested that early studies, such as Aggarwal (1981) and Roll (1992) may have produced spurious results by not testing the variables for stationarity and cointegration. To address this issue, Bahmani-Oskooee and Sohrabian (1992) employed monthly data for the period 1973–1988 on the S&P500 index and the US dollar effective exchange rate. They reported bidirectional causality between the two variables in the short-run, but no cointegration and hence no long-run relationship between exchange rates and stock prices. Subsequent studies for other countries, such as Morley and Pentecost (2000), Nieh and Lee (2001), Yang and Doong (2004); Kollias et al. (2012), Wong (2017), Salisu and Ndako (2018) also find little support for cointegration. Nevertheless, researchers such as Liang et al. (2013) and Tang and Yao (2018) have provided evidence for a long-run relationship between exchange rates and stock prices in some countries.

As pointed out in Bahmani-Oskooee and Saha (2015, 2018), even results from several later studies should be reexamined if they assume a linear and symmetric relationship between stock prices and exchange rates. The relationship between the two variables is likely to be asymmetric and nonlinear. For example, an appreciating currency may reduce the cost of imported inputs, leading to a positive relationship between exchange rates and stock prices rather than the expected negative relationship. Similarly, currency depreciations that increase profits for export-oriented firms can increase the costs of imported inputs. Some producers will increase prices, but others may be forced to absorb the cost increases. Bahmani-Oskooee and Saha (2015) suggest that stock prices will show a larger response to currency appreciations than to depreciations. Similarly, Salisu and Ndako (2018) note that bad news has a greater impact on investor sentiment in stock markets than good news. Hence, falling stock prices after bad news may have a larger effect on the exchange rate than rising stock prices following good news. Furthermore, adjustment to long-run equilibrium may exhibit nonlinearities due to the presence of fixed adjustment costs, transaction costs, or policy interventions, such as exchange rate management or commodity price stabilization (Balke and Fomby, 1997).

The purpose of this paper is to examine the links between stock prices and exchange rates for the G7 countries (Canada, France, Germany, Italy, Japan, the UK, and the US). Our primary objective is to empirically test two competing theories—the flow-oriented approach that exchange rates affect stock prices versus the portfolio balance approach that stock prices affect exchange rates. We first examine the short-run and long-run relationship between stock prices and exchange rates using the linear autoregressive distributed lag (ARDL) model of Pesaran et al. (2001). Then, we adopt the nonlinear ARDL (NARDL) model of Shin et al. (2013) and allow for both short-run and long-run asymmetries in the relationship between stock prices and exchange rates via a partial sum decomposition of the explanatory variables. We also test for and include structural breaks in our analysis. The asymmetries are introduced by differentiating between currency appreciations and depreciations and positive and negative changes in stock prices. Furthermore, a vector autoregressive (VAR) model is estimated to test for cointegration between the variables using the Johansen (1998) procedure. These results are used to determine the direction of causality between the variables based upon the Granger (1969) causality test. Monthly data from mid-1973 to February 2020 are collected for the USA and Japan, while the time span begins in 1979 for Canada, 1984 for the UK, 1987 for Germany, 1990 for France, and December of 1997 for Italy. The data extend to February 2020 for all countries except Italy, for which our data ends in February of 2018.

We find short-run support for both the flow-oriented and the portfolio balance approaches using both linear and nonlinear models. The linear ARDL model does not support either theory in the long-run and the nonlinear ARDL model fails to support the flow-oriented approach. However, the nonlinear asymmetric ARDL model indicates a link from stock prices to exchange rates that is non-spurious in four of the G7 countries. Moreover, Granger causality tests provide evidence of unidirectional causality from stock prices to exchange rates in six of the G7 countries. By allowing for structural breaks and using a longer timeframe and a data set that includes more observations after the Global Financial Crisis than previous studies, we provide stronger long-run support for the portfolio balance approach than is found in recent studies, such as Bahmani-Oskooee and Saha (2018). Also, even though panel techniques have been criticized in the literature for producing outcomes that are sometimes driven by a few variables or countries, we generally confirm recent panel data results provided by Salisu and Ndako (2018) and Xie et al. (2020) that validate the portfolio balance approach. We find moderate long-support for the portfolio balance approach in some G7 countries, but we find no long-run evidence for the flow-oriented approach.

[2] Alternatively, researchers may discover either bi-directional causality, or no causality. A finding of no causality is consistent with Frenkel (1976) and the subsequent monetary models of exchange rate determination. They suggest that both exchange rates and stock prices are affected by some common variable, such as interest rates.

## 2. Literature review

Since Bahmani-Oskooee and Saha (2015) provide an extensive literature review of the empirical tests of the flow-oriented and stock-oriented models of exchange rate determination, this section focuses primarily on recent articles and research specific to the G7 countries.[3] Except for China, these are the world's largest economies, and it is surprising that there are only a limited number of studies of exchange rate determination for these countries.

In an early study that included the G7 countries among a group of 27 countries, Roll (1992) found a positive relationship between exchange rates and stock prices using daily data over the period 1988–1991. However, as mentioned by Bahmani-Oskooee and Sohrabian (1992), many early studies did not test the variables for unit roots and cointegration. Using monthly data on the S&P500 index and the US dollar effective exchange rate for the period 1973–1988, Bahmani-Oskooee and Sohrabian (1992) discovered that the variables were not stationary and not cointegrated, but they find short-run bidirectional Granger causality between stock prices and exchanges rates. Subsequent research, such as Morley and Pentecost (2000), Nieh and Lee (2001), and Yang and Doong (2004), have reported similar results for various G7 countries. That is, much of the literature suggests that the two variables are related in the short-run, but there is only limited evidence of a long-run relationship between exchange rates and stock prices.

In contrast, Chow et al. (1997) obtained noticeably different results using monthly data over the period 1977–1989. They found no relationship between US excess stock returns and real exchange rate returns over short horizons, but a positive relationship over horizons of longer than six months. Similarly, Ajayi and Mougoue (1996) employed Engle-Granger cointegration procedure and causality tests with daily data for the period 1985–1991 for the G7 countries plus the Netherlands and found cointegration in all cases and feedback between the two markets in both the short-run and on the long-run. However, results varied considerably between countries. Rising stock prices led to short-run currency depreciation in all countries, but long-run currency appreciation only in Germany, the UK, and the US. Currency depreciation leads to falling stock prices in the short-run in five of eight countries and in all eight countries in the long-run. Similarly, Chen and Chen (2012) employed both linear and non-linear Granger causality tests for 12 OECD countries (that included the G7 countries) using monthly data for the period 1993–2007, and they also reported considerable variation across countries. For example, linear causality tests suggested bidirectional causality between stock prices and exchange rates for Canada, Germany, Italy, the US, and the UK, but no evidence of causality in any direction for France and Japan. When using nonlinear causality tests, they found evidence of unidirectional causality from exchange rates to stock prices for Germany only, and bidirectional causality for the other G7 countries.

Using rolling cointegration analysis with daily data for the period 2002–2008 on the Euro-US dollar exchange rate and the FTSE Eurotop 300 and FTSE eTX All-Share Index, Kollias et al. (2012) found no evidence of cointegration. However, their rolling causality tests showed a time-varying causality that depends on market conditions. They found that exchange rate movement Granger-causes stock price changes under normal market conditions, but that causality is reversed under stressful market conditions. Since causality relations may vary over time, Caporale et al. (2014) employed the Gregory and Hansen (1996) cointegration test allowing for structural breaks in weekly data over the period 2003–2011 for the US, UK, Canada, Japan, Switzerland, and the Euro area. They found evidence of cointegration in all cases. Their bivariate UEDCC-GARCH model showed evidence of unidirectional causality from stock returns to exchange rate changes in the US and the UK, a unidirectional causality from exchange rate changes to stock returns in Canada, and bidirectional causality in the Euro area and Switzerland. Finally, using cointegration and error correction models with an asymmetric threshold approach, Koulakiotis et al. (2015) found a causal relationship from stock prices to exchange rates in Canada, the US, and the UK based on monthly data for the period 1990–2014. Using a threshold cointegration methodology with monthly data over the period 2000–2014 for eight European countries that included France, Germany, Italy, and the UK, Kollias et al. (2016) found evidence of cointegration in all countries. Their results showed unidirectional causality from stock prices to exchange rates for all countries, except for the UK, where bidirectional causality was found. Wong (2017) used an MGARCH model with quarterly data for the period 1985–2015 for seven countries that included Japan, the UK, and Germany and found a significant and negative relationship between stock prices and exchange rates for the UK, but an insignificant relationship for Japan and Germany. Although most of these studies have not adopted ARDL models used in this paper, they suggest causality may vary between countries within the G7, that relationships may vary over time, and that findings may differ between linear, symmetric models and nonlinear, asymmetric models.

Four recent studies that are particularly relevant for our research are now discussed. Salisu and Ndako (2018) employed panel linear and non-linear ARDL models with monthly data for the period 2004–2017 for 32 OECD countries that included Canada, France, Germany, Italy, Japan, and the UK. They found evidence supporting the portfolio balance approach. Interestingly, results were stronger for the Euro area than for the non-Euro area, and stronger across all countries after September of 2007, or in the aftermath of the Global Financial Crisis. Similarly, Xie et al. (2020) examined a panel of 26 countries that included the G7 countries with daily data for the period 1998–2019. Using symmetric and asymmetric panel Granger causality tests, they found evidence of causality from stock prices to exchange rates, in support of the portfolio balance approach. However, they found no evidence for the flow-oriented approach involving causality from exchange rates to stock prices.

Our research builds upon two studies by Bahmani-Oskooee and Saha (2016, 2018). First, Bahmani-Oskooee and Saha (2016) employed both linear and nonlinear ARDL models with monthly data for nine countries that included Canada, Japan, and the UK. While both models showed an insignificant long-run relationship and no evidence of cointegration, both the linear and nonlinear

[3] Bahmani-Oskooee and Saha (2015) summarize the results of the various univariate and multivariate studies for developed and developing countries in North America, Europe, Asia, the Middle East and Africa and illustrate how to test for asymmetric relationships between exchange rates and stock prices using monthly US data for the period 1973–2014.

models (with positive and negative exchange rates changes) showed significant short-run effects from exchange rates to stock prices. Second, Bahmani-Oskooee and Saha (2018) employed linear and nonlinear ARDL models with monthly data for 24 countries that included all the G7 countries, except Italy. Adopting the linear ARDL model with the stock price as the dependent variable, the authors found no evidence of cointegration in any G7 country and with the nonlinear ARDL model they found evidence of asymmetric cointegration for only Canada. With the exchange rate as the dependent variable, they found significant short-run effects for all countries that lasted into the long-run for Japan based on the linear ARDL model and for Germany based on the nonlinear ARDL model. Although we adopt the same methodology as Bahmani-Oskooee and Saha (2016, 2018), we focus specifically upon univariate models for the G7 countries. Also, we employ a longer data series for each of the G7 countries and we use more recent data to obtain more observations for the post financial crisis period. By examining more recent data and allowing for structural breaks, we provide stronger long-run support for the portfolio balance approach than discovered in previous research.

## 3. Data and methodology

### 3.1. Data

To examine the links and interactions between stock prices and exchange rates for the G7 countries, monthly data on stock price indices and nominal effective exchange rates are used. Sample period varies from one country to another based on data availability (see Appendix for data description). End-of-month closing prices for the seven national stock indexes are extracted from Yahoo Finance. Nominal effective exchange rates (NEERs) are extracted from the Bank of International Settlements. We use the effective exchange rate because the G7 countries trade with many trading partners; thus, using the effective exchange rate gives better results than the bilateral exchange rate.

### 3.2. Methodology

The methodology adopted in this paper to examine the links and interactions between stock prices and exchange rates is based on previous literature (see, for example, Bahmani-Oskooee and Saha, 2018; Chen and Chen, 2012; Bahmani-Oskooee and Sohrabian, 1992). The following long-run relationship between the two variables is used:

$$p_t = \alpha_0 + \alpha_1 ex_t + \nu_t \tag{1}$$

where $p_t$ and $ex_t$ is the aggregate stock price index and the effective nominal exchange rate at time$t$, respectively, and $\nu_t$ is a random error term. The variables are in logarithmic form. An increase in the effective exchange rate implies appreciation in the currency of the country under study (G7 country).

Equation (1) provides a test for the flow-oriented model (goods market approach) that exchange rates affect stock prices. This equation is a long-run relationship between the two variables. Therefore, the estimate $\alpha_1$ captures only the long-run effect of exchange rate changes on the stock price. To allow for short-run effects, equation (1) can be converted into an error-correction model. However, the ARDL of Pesaran et al. (2001) is a convenient approach in this case because it can estimate both short-run and long-run effects in one step. Another convenient property of the ARDL model is that it can be applied regardless of whether the variables are integrated of order one, zero, or a combination of both (Pesaran et al., 2001). As in Pesaran et al. (2001), the following conditional error correction model (ECM) can be written:

$$\Delta p_t = \beta_0 + \beta_1 DV_t + \beta_2 p_{t-1} + \beta_3 ex_{t-1} + \sum_{j=1}^{m} \omega_j \Delta p_{t-j} + \sum_{j=0}^{n} \vartheta_j \Delta ex_{t-j} + \epsilon_t \tag{2}$$

where $\beta_2$ and $\beta_3$ are the long-run coefficients, $\omega_j$ and $\vartheta_j$ are the short-run coefficients, and $m$ and $n$ are the optimal lags on the first-differenced variables selected by AIC. $DV_t = (DV_{1t}, \cdots, DV_{jt})^{'}$ is a vector of $j$ dummy variables, where $DV_{jt} = 1$ if observation $t$ belongs to the $j^{th}$ period and 0, otherwise. The time periods of $DV_t$ will be defined in the next section. To see if the variables are cointegrated, equation (2) is estimated by the ordinary least squares (OLS) method, and then the $F_{PSS}$ test statistic of Pesaran et al. (2001) or the $t_{BDM}$-test statistic of Banerjee et al. (1998) are used to establish cointegration. $F_{PSS}$ tests the null hypothesis of no-cointegration $(\beta_2 = \beta_3 = 0)$ against the alternative of cointegration$(\beta_2 \neq 0, \beta_3 \neq 0)$, whereas $t_{BDM}$ tests the null hypothesis of no-cointegration $(\beta_2 = 0)$ against the alternative of cointegration$(\beta_2 < 0)$..

The linear ARDL model of Pesaran et al. (2001) assumes that stock prices respond symmetrically to changes in exchange rates. However, this assumption may not be realistic given our discussion about asymmetries in the relationship between stock prices and exchange rates. To allow for asymmetries in the relationship, we employ the NARDL model of Shin et al. (2013), which is an asymmetric expansion of the linear ARDL model of Pesaran et al. (2001).

Shin et al. (2013) show that the following asymmetric long-run relationship can be written:

$$p_t = \alpha_0 + \delta^+ ex_t^+ + \delta^- ex_t^- + \varepsilon_t \tag{3}$$

where $ex_t^+$ and $ex_t^-$ are partial sum processes of positive and negative changes in$ex_t$, calculated as:

$$ex_t^+ = \sum_{i=1}^{t} \Delta ex_i^+ = \sum_{j=1}^{t} max(\Delta ex_i, 0) \quad (4)$$

$$ex_t^- = \sum_{i=1}^{t} \Delta ex_i^- = \sum_{i=1}^{t} min(\Delta ex_i, 0) \quad (5)$$

where $\Delta ex^+$ and $\Delta ex^-$ are positive and negative changes in the exchange rate, reflecting domestic currency appreciation and depreciation, respectively. Thus, $\delta^+$ and $\delta^-$ represent long-run parameters capturing the effects domestic currency appreciations and depreciations, respectively.

Then, following Shin et al. (2013), the NARDL model is obtained by replacing $ex_t$ in the linear ARDL model in equation (2) by $ex_t^+$ and$ex_t^-$:

$$\Delta p_t = \gamma_0 + \gamma_1 DV_t + \rho_0 p_{t-1} + \eta^+ ex_{t-1}^+ + \eta^- ex_{t-1}^- + \sum_{j=1}^{k} \gamma_{1j} \Delta p_{t-j} + \sum_{j=0}^{l} \gamma_{2j}^+ \Delta ex_{t-j}^+ + \sum_{j=0}^{m} \gamma_{3j}^- \Delta ex_{t-j}^- + \varepsilon_t \quad (6)$$

Due to the way the partial sum processes ($ex_t^+$ and $ex_t^-$) are generated, Shin et al. (2013) call equation (6) a "nonlinear ARDL (NARDL) model." In their NARDL model, nonlinearities in both the short-run and long-run, are introduced "via positive and negative partial sum decompositions of the explanatory variables." Since the NARDL model (6) is an asymmetric extension of the linear ARDL model, the estimation procedure of the NARDL is like that of the linear ARDL model. Thus, the NARDL model (6) is estimated by the OLS method, and asymmetric (nonlinear) cointegration is established using the $F_{PSS}$ statistic or the $t_{BDM}$ statistic. $F_{PSS}$ tests the null hypothesis of asymmetric no-cointegration ($\rho_0 = \eta^+ = \eta^- = 0$) against the alternative of cointegration, and $t_{BDM}$ tests the null hypothesis of no-cointegration ($\rho_0 = 0$) against the alternative of cointegration($\rho_0 < 0$). The next step is to test for long-run and short-run asymmetries. Long-run asymmetry is tested by evaluating the null hypothesis of symmetry$\delta^+ = \delta^-$, where $\delta^+ = -\eta^+/\rho_0$ and$\delta^- = -\eta^-/\rho_0$. Rejecting this null indicates the existence of a long-run asymmetric effect. Short-run additive (or impact) asymmetry is tested by evaluating the null hypothesis of symmetry$\sum_{j=0}^{p} \gamma_{2j}^+ = \sum_{j=0}^{n} \gamma_{3j}^-$. Rejecting this null implies that there is a short-run asymmetric effect. In addition to short-run impact asymmetry, short-run adjustment asymmetry can be observed if the variables $\Delta ex_t^+$ and $\Delta ex_t^-$ take different lags. Also, short-run asymmetry can be established if $\gamma_{2j}^+ \neq \gamma_{3j}^-$ for each individual $j$ (Bahmani-Oskooee and Mohammadian, 2018).

To examine the stock-oriented model (portfolio balance approach), the dependent and independent variables are switched in equations (1) - (6). This gives us the following linear and nonlinear ARDL models:

$$\Delta ex_t = \alpha_0 + \alpha_1 DV_t + \alpha_2 e_{xt-1} + \alpha_3 p_{t-1} + \sum_{j=1}^{m} \omega_j \Delta ex_{t-j} + \sum_{j=0}^{n} \vartheta_j \Delta p_{t-j} + \upsilon_t \quad (7)$$

$$\Delta ex_t = \gamma_0 + \gamma_1 DV_t + \rho_0 e_{t-1} + \eta^+ p_{t-1}^+ + \eta^- p_{t-1}^- + \sum_{j=1}^{k} \gamma_{1j} \Delta ex_{t-j} + \sum_{j=0}^{l} \gamma_{2j}^+ \Delta p_{t-j}^+ + \sum_{j=0}^{k} \gamma_{3j}^- \Delta p_{t-j}^- + \varepsilon_t \quad (8)$$

where and then define $p_t^+$ and $p_t^-$ as partial sum processes of positive and negative changes in stock price index$p_t$.

## 4. Empirical results

### 4.1. Preliminary analysis

To examine any possible co-movement between the exchange rate and the stock price of each country, Fig. 1 plots the two variables for each country over the sample period. The solid line (left scale) is the logarithm of the nominal effective exchange rate, and the dashed line (right scale) is the logarithm of the stock price index. A visual inspection suggests that there is some degree of co-movement between the two variables, although this degree of co-movement differs across the G7 countries. To further examine the degree of correlation between the variables, Fig. 1 also reports the correlation coefficient between the two variables for each country over the entire sample period and for different sub-periods. Over the entire sample period, the correlation coefficient varies between −0.55 for the US to +0.64 for Japan. The correlation coefficient is negative for Canada, Italy, the UK and the US, ranging from −0.02 for Canada to −0.55 for the US., and positive for France, Germany and Japan, ranging from 0.13 in France to 0.64 in Japan. Given the definition of the exchange rate, the positive correlation coefficients for France, Germany and Japan imply that currency appreciations are

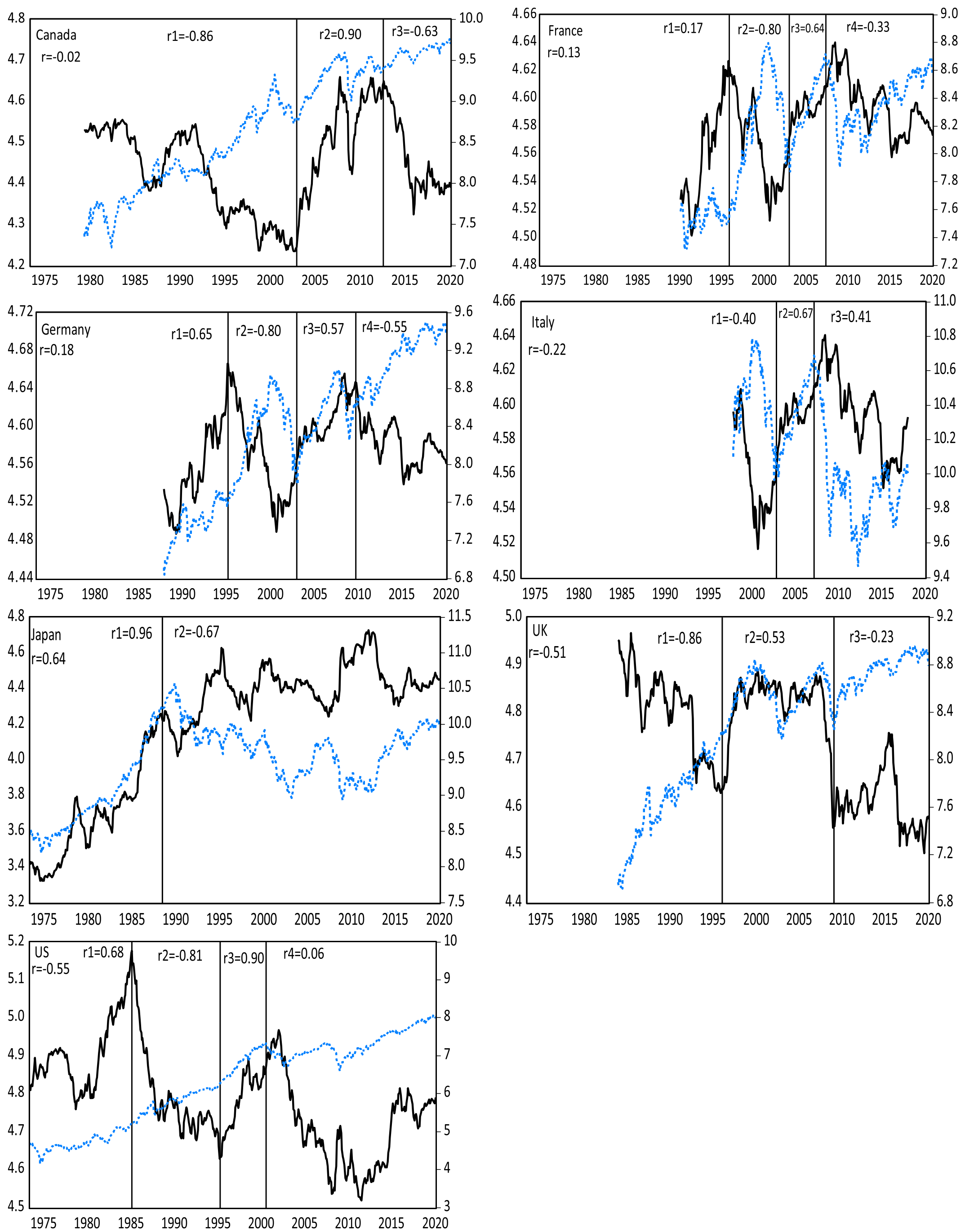


**Fig. 1.** Evolution of stock prices and nominal effective exchange rates (in logarithms). Notes: solid black line (left scale) is the nominal effective exchange rate. Dashed blue line (right scale) is the stock price index. Variables are in logarithmic form. r stands for the correlation coefficient over the entire sample, and r1, r2, r3, and r4 stand for the correlation coefficient over the different sub-periods. (For interpretation of the references to colour in this figure legend, the reader is referred to the web version of this article.)

associated with rising stock prices, while the negative correlation coefficients for Canada, Italy, the UK and the US. imply that currency appreciations are associated with falling stock prices.

An examination of results over time shows that the correlation coefficient has not been stable and that it changes in sign and/or size between various sub-periods across all the G7 countries.[4] For example, the correlation coefficient in Canada is negative, but near zero (−0.02) for the entire sample period. However, it is −0.86 during the first sub-period (1979:06 – 2003:01), then it reverses to +0.90 during the second sub-period (2003:02 – 2012:09), and changes again to −0.63 during the third sub-period (2012:1 – 2020:02). In the US., the correlation coefficient is −0.55 over the entire sample period, but +0.68 during the first sub-period (1973.:06 – 1983:03), −0.81 in the second sub-period (1985:04 – 1995:05), +0.90 during the third sub-period (1995:06 – 2000:08), and it drops to +0.06 during the fourth sub-period (2000:09 – 2020:02). A similar picture appears in the rest of the countries, suggesting that the relationship between exchange rates and stock market prices may be quite complex and that a simple linear model may not be able to adequately capture this relationship.

This preliminary analysis using correlation coefficients shows the degree of co-movement between stock prices and exchange rates, but it does not provide information about the short-run or the long-run relationships, nor does it identify causation in any direction. To further explore the relationships between variables, we first check for stationarity of variables with unit root tests. We then utilize both linear and nonlinear ARDL models to examine the short-run and long-run relationships between stock prices and exchange rates. We also estimate a vector autoregressive (VAR) model to examine the robustness of the ARDL results.

### 4.2. Unit root tests

While both the linear and nonlinear ARDL models can handle variables that are $I(1)$ or $I(0)$ or a combination of both, they cannot handle variables that are$I(2)$. In addition, before estimating a VAR model and testing for Granger causality, it is necessary to investigate the stationarity and cointegration properties of the variables, because finding evidence of cointegration affects the model specification used to test for causality. Specifically, if the variables are not cointegrated, a VAR model is specified in first differences, but if cointegrated, a vector error correction model (VEC) is used to examine Granger causality (Engle and Granger, 1987). Therefore, determining the order of integration of the variables is important before proceeding to the next steps. We do this by using the Augmented Dickey and Fuller (1979, ADF hereafter) test, Phillips and Perron (1988, PP hereafter) test, and the KPSS test by Kwiatkowski *et al.* (1992). Whereas the ADF and PP tests test the null hypothesis of a unit root, the KPSS test tests the null hypothesis of stationary. The results are reported in Table 1. The three tests suggest that all the variables are nonstationary in levels, but stationary in their first differences, which rules out the possibility of $I(2)$ variables.

Given the long and volatile sample period, testing the variables for structural breaks is important. We do this by using the multiple structural breaks test developed by Bai and Perron (1998, 2003). We allow for a maximum of five breaks and testing the null hypothesis of $l+1$ vs. $l$ sequentially determined structural breaks. The results reported Table 2 suggest two to five structural breaks in the variables. Although the dates of the breaks are not uniformly distributed across the G7 countries, they generally correspond with major events that happened over the sample period. Accordingly, and to avoid dealing with numerous dummy variables corresponding to every structural break for each country, we construct dummy variables associated with five major world-wide events that reflect most of the individual country structural breaks.[5] The dummy variables included in the ARDL and NARDL models are: $DV_{85}$ for the period 1985:09–1986:07 to capture the effect of the Plaza Accord, $DV_{97}$ for the period 1997:06–1998:12 to capture the effect of the Asian financial crisis,$DV_{03}$ for the period 2003:03–2003:12 to capture the Second Gulf War, $DV_{07}$ for the period 2007:06–2009:12 to capture the effect of the US financial crisis, and $DV_{13}$ for the period 2013:01–2014:03 to capture the peak in oil prices in 2012 and the sharp oil price decreases that followed. Each of the dummy variables are equal to one in the specified period, and zero in all other periods.

### 4.3. The flow-oriented model

In this section, we adopt the flow-oriented model (or the goods market approach) to examine the impact of exchange rate changes on stock prices and to determine whether the effects are symmetric or asymmetric. To that end, the linear and nonlinear ARDL models in equations (1) – (6) are estimated by the OLS method. Allowing for a maximum of 12 lags, the optimal number of lags is determined by AIC. Table 3 reports the results from estimating the linear ARDL model. The table contains short-run (in Panel A) and long-run estimates (in Panel B) and some diagnostic checks (in Panel C). Focusing on the short-run effects of exchange rate changes on stock prices, Panel A shows that there is at least one statistically significant coefficient in all the countries, except in Italy. This indicates that changes in exchange rates do have short-run effects on the stock prices of the G7 countries, except Italy. The effect is negative in France, Germany, Japan, the US and the UK, implying that currency depreciations in these countries increases their stock prices. This result

[4] We are not attempting to explain why results vary so dramatically over time, nor do we try to identify the types of events that could have been responsible. The choice of sub-periods varies by country and is obtained through a visual inspection of the graph of each country. Nevertheless, many of the breaks between sub-periods, do correspond to major events that could have altered the relationship between stock prices and exchange rates. For example, the first sub-period for Japan (1973:06 – 1988:07), with a correlation coefficient of +0.96, occurs during the period of rapidly rising Japanese stock prices that lasted from mid-1950 to 1989. The second sub-period (1988:08 – 2020:02) with a correlation coefficient of −0.67 reflects the aftermath of the stock market crash of 1989 and the drastic downward trend in stock prices that lasted until 2013.

[5] To check the robustness of the results, we estimated the ARDL and NARDL models using the specific individual months as suggested by Bai-Perron test and the results did not significantly change. All unreported results are available upon request from the authors.

**Table 1**
Unit root tests.

| Country | Variable | ADF test | | | PP test | | | KPSS test | | |
|---|---|---|---|---|---|---|---|---|---|---|
| | | Level | | First diff. | Level | | First diff. | Level | | First diff. |
| | | Trend | No trend | No trend | Trend | No trend | No trend | Trend | No trend | No trend |
| Canada | *p* | −3.49(1)** | −1.48(1) | −19.61(0)* | −3.49** | −1.38 | −19.62* | 0.276* | 2.752* | 0.056 |
| | *ex* | −1.89(1) | −1.89(1) | −17.06(0)* | −1.80 | −1.80 | −16.96* | 0.327* | 0.325 | 0.093 |
| France | *p* | −1.88(0) | −1.53(0) | −17.49(0)* | −2.04 | −1.61 | −17.49* | 0.274* | 1.322* | 0.072 |
| | *ex* | −2.58(3) | −2.73(3)*** | −9.47(2)* | −2.33 | −2.49 | −15.22 | 0.151** | 0.515** | 0.118 |
| Germany | *p* | −2.49(0) | −1.65(0) | −18.67(0)* | −2.65 | −1.65 | −18.69* | 0.192** | 2.087* | 0.095 |
| | *ex* | −2.26(2) | −2.39(2) | −12.77(1)* | −1.75 | −2.13 | −14.03* | 0.155** | 0.358*** | 0.096 |
| Italy | *p* | −2.58(4) | −1.84(4) | −8.63(2)* | −2.81 | −1.64 | −14.89* | 0.118 | 1.214* | 0.094 |
| | *ex* | −1.96(1) | −1.88(1) | −11.54(0)* | −1.81 | −1.69 | −11.45* | 0.303* | 0.401*** | 0.099 |
| Japan | *p* | −1.62(0) | −1.74(0) | −−22.61(0)* | −1.70 | −1.79 | −22.64* | 0.458* | 0.915* | 0.167 |
| | *ex* | −1.84(16) | −2.21(16) | −6.13(15)* | −1.59 | −1.78 | −16.87* | 0.598* | 2.407* | 0.184 |
| UK | *p* | −2.32(0) | −2.53(0) | −15.51(1)* | −2.28 | −2.53 | −20.37* | 0.470* | 2.185* | 0.321 |
| | *ex* | −2.57(1) | −1.98(1) | −16.31(0)* | −2.47 | −1.89 | −16.26* | 0.239* | 1.373* | 0.051 |
| US | *p* | −2.04(0) | −0.35(0) | −22.87(0)* | −2.16 | −0.37 | −22.89* | 0.448* | 2.927* | 0.072 |
| | *ex* | −2.25(3) | −1.91(3) | −14.81(1)* | −2.15 | −1.82 | −16.79* | 0.110 | 1.287* | 0.073 |

*, **, *** denotes rejection of the null hypothesis at the 1%, 5% and 10% significance level. The number of lags (in parentheses) is selected by minimizing Akaike Information Criterion (AIC). The bandwidth for PP and KPSS tests is selected automatically by Newey-West Bandwidth, using Barlett Kernel spectral estimation method. *p* and *ex* are stock price index and nominal effective exchange rate, respectively. Variables are in logarithmic form. The 1%, 5% and 10% critical values for the ADF and PP tests are −3.98, −3.42, and −3.13 for the test with a trend, and −3.44, −2.87, and −2.57 for the test with no trend. The 1%, 5% and 10% critical values for the KPSS test are 0.216, 0.146, and 0.119 for the test with a trend, and 0.739, 0.463, and 0.347 for the test with no trend.

**Table 2**
Bai-Perron multiple breakpoint tests of $L + 1$ vs. $L$ sequentially determined breaks.

| Country | Variable | 0 vs. 1 | 1 vs. 2 | 2 vs. 3 | 3 vs. 4 | 4 vs. 5 | No. of breaks | Dates of breaks |
|---|---|---|---|---|---|---|---|---|
| Canada | *p* | 1954.70** | 900.58** | 524.63** | 238.51** | 0.00 | 4 | 1985 M11, 1996 M09, 2005 M02, 2013 M12 |
| | *ex* | 115.62** | 830.42** | 301.52** | 131.54** | 0.00 | 4 | 1985 M05, 1993 M08, 2005 M07, 2014 M02 |
| France | *p* | 1500.49** | 117.17** | 57.10** | 8.95 | | 3 | 1997 M12, 2008 M09, 2013 M10 |
| | *ex* | 117.55** | 183.34** | 33.63** | 29.04** | 0.00 | 4 | 1994 M09, 1999 M03, 2007 M11, 2012 M05 |
| Germany | *p* | 1449.18** | 793.30** | 95.14** | 0.000 | | 3 | 1996 M11, 2005 M12, 2013 M09 |
| | *ex* | 259.89** | 61.38** | 125.86** | 216.21** | 27.82** | 5 | 1992 M10, 1997 M08, 2003 M05, 2010 M03, 2015 M01 |
| Italy | *p* | 635.16** | 32.20** | 112.51** | 20.42** | 0.00 | 4 | 2001 M09, 2005 M01, 2008 M09, 2014 M01 |
| | *ex* | 108.27** | 251.81** | 38.90** | 123.03** | 0.00 | 4 | 2003 M04, 2007 M03, 2010 M05, 2015 M01 |
| Japan | *p* | 1320.12** | 141.94** | 359.50** | 9.84 | | 3 | 1984 M10, 2000 M10, 2013 M03 |
| | *ex* | 2557.69** | 894.15** | 61.21** | 30.03** | 0.00 | 4 | 1985 M10, 1993 M02, 2001 M10, 2008 M10 |
| UK | *p* | 1096.98** | 424.84** | 242.76** | 34.57** | 0.00 | 4 | 1989 M06, 1995 M11, 2005 M07, 2013 M01 |
| | *ex* | 988.40** | 77.70** | 130.06** | 21.54** | 1.96 | 4 | 1991 M11, 1997 M06, 2003 M02, 2008 M09 |
| US | *p* | 2095.02** | 1273.34** | 717.12** | 50.11** | 0.00 | 4 | 1985 M11, 1995 M11, 2003 M12, 2013 M03 |
| | *ex* | 429.25** | 181.32** | 228.40** | 127.19** | 108.05** | 5 | 1980 M09, 1987 M09, 1997 M08, 2004 M09, 2013 M03 |

** denotes significance at the 5% level. The 5 percent critical values for the hypotheses 0 vs. 1, 1 vs. 2, 2 vs. 3, 3 vs. 4, and 4 vs. 5 breaks are 8.58, 10.13, 11.14, 11.83, 12.25. The numbers in the cell corresponding to the null hypothesis being tested are the scaled *F-statistic* values. The test is implemented allowing for heterogeneous error distributions across breaks.

suggests that these countries have more export-oriented firms. Therefore, currency depreciations improve their international competitiveness, which results in more exports, and this leads to higher expected cash flows and higher stock prices. The effect is negative for Canada, suggesting that Canadian dollar depreciation causes Canada's stock prices to decrease.

Since we found that changes in exchange rates have significant short-run effects on stock prices in the G7, except in Italy, we next examine the cointegration results to see if short-run effects continue into the long-run. At the 5 percent significance level or lower, the $F_{PSS}$ and $t_{BDM}$ statistics in Panel C suggest lack of cointegration in all the countries. However, at the 10 percent level, the $F_{PSS}$ ($t_{BDM}$) statistic shows evidence of cointegration in the UK (Canada). Alternatively, cointegration can be established by examining the significance of the error correction term($ECM_{t-1}$). A negative and significant $ECM_{t-1}$ provides evidence of cointegration. The results in Panel C show that this term is negative and significant for Canada and France at the 10 percent significance level, and for the UK at the 5 percent level. Thus, our results suggest either lack of, or poor evidence for cointegration. In addition, Panel B shows that none of the exchange rates has a statistically significant long-run coefficient estimate.

Overall, our results of a significant short-run relationship between exchange rates and stock prices, lack of cointegration, and an

**Table 3**
Linear ARDL model coefficient estimates and diagnostic checks (flow-oriented model).

| A. Short-run estimates | | | | | | | |
|---|---|---|---|---|---|---|---|
| | Canada | France | Germany | Italy | Japan | UK | U. S |
| Constant | 0.363(2.984)* | −0.873(1.809)*** | 0.014(0.035) | 0.815(1.111) | 0.085(2.076)** | 0.235(1.916)*** | 0.172(1.864)** |
| *Trend* | 0.0002(3.242)* | N.S | 0.0001(1.856)*** | −0.0002(1.750)*** | N.S | N.S | 0.0002(2.970)* |
| $DV_{85}$ | N.S | N.S | N.S | N.S | 0.036(2.089)** | N.S | N.S |
| $DV_{97}$ | N.S | N.S | N.S | N.S | N.S | N.S | 0.027(2.529)** |
| $DV_{03}$ | N.S | N.S | N.S | N.S | N.S | N.S | N.S |
| $DV_{07}$ | N.S | −0.026(2.325)** | N.S | N.S | −0.021(2.070)** | N.S | −0.024(2.864)* |
| $DV_{13}$ | N.S | N.S | N.S | N.S | −0.012(2.052)** | N.S | N.S |
| $p_{t-1}$ | −0.043(3.519)* | −0.016(1.604) | −0.030(2.433)** | −0.048(2.593)** | 0.007(0.848) | −0.013(2.752)* | −0.023(3.083)* |
| $ex_{t-1}$ | −0.010(0.539) | 0.212(2.011)** | 0.410(0.496) | −0.058(0.385) | | −0.027(1.272) | −0.013(0.745) |
| $\Delta p_{t-1}$ | 0.059(1.266) | | | 0.070(1.079) | | | |
| $\Delta p_{t-2}$ | | | | −0.047(0.730) | | | |
| $\Delta p_{t-3}$ | | | | 0.127(1.986)** | | | |
| $\Delta p_{t-4}$ | | | | 0.127(2.000)** | | | |
| $\Delta ex$ | 0.761(5.252)* | −0.862(1.704)*** | −0.873(2.002)** | | −0.393(3.748)* | −0.248(1.900)*** | −0.520(4.345)* |
| $\Delta ex_{t-1}$ | | | | | 0.185(1.705)*** | | |
| $\Delta ex_{t-2}$ | | | | | −0.215(1.973)** | | |
| $\Delta ex_{t-3}$ | | | | | 0.085(0.781) | | |
| $\Delta ex_{t-4}$ | | | | | −0.089(0.818) | | |
| $\Delta ex_{t-5}$ | | | | | 0.250(2.413)** | | |
| B. Long-run estimates | | | | | | | |
| *constant* | 8.390(4.499)* | −75.470(1.171) | 0.471(0.035) | 17.030(1.185) | 7.316(3.111)* | 18.560(2.749)* | 7.410(1.849)*** |
| *Trend* | 0.005(14.004)* | N.S | 0.005(4.674)* | −0.003(2.579)** | N.S | N.S | 0.007(11.295)* |
| $ex_t$ | −0.223(0.536) | 18.353(13.00) | 1.388(0.465) | −1.214(0.384) | 0.586(1.064) | −2.091(1.470) | −0.578(0.712) |
| C. Diagnostic checks | | | | | | | |
| $F_{PSS}$ | 4.41 | 2.84 | 2.64 | 2.27 | 2.23 | 3.74*** | 3.61 |
| $t_{BDM}$ | −0.043(3.519)*** | −0.016(1.604) | −0.030(2.433) | −0.048(2.593) | −0.021(2.070) | −0.013(2.752) | −0.023(3.083) |
| $ECM_{t-l}$ | −0.043(3.644)*** | −0.012(2.928)*** | −0.030(2.821) | −0.048(2.620) | −0.012(2.592) | −0.013(3.357)** | −0.023(3.301) |
| *LMtest* | 0.258[0.6117] | 0.588[0.4438] | 0.734[0.3922] | 0.121[0.7286] | 0.147[0.7017] | 0.052[0.8193] | 0.062[0.8035] |
| *RESETtest* | 0.442[0.5064] | 0.012[0.9129] | 0.153[0.6961] | 1.127[0.2896] | 1.369[0.2426] | 0.132[0.7164] | 3.096[0.07790]*** |
| $\overline{R}^2$ | 0.086 | 0.031 | 0.024 | 0.039 | 0.049 | 0.024 | 0.054 |
| *CUSUM* | Stable | Stable | Stable | Stable | Stable | Stable | Stable |
| *CUSUMQ* | Unstable | Unstable | Unstable | Stable | Unstable | Unstable | Unstable |

| Critical values | | | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|
| Significance level | $F_{PSS}$ for $K = 1$ and $n = 1000$: upper values | | | | $t_{BDM}$ for $K = 1$ and $n = 500$ | | *t-value* for significance of $ECM_{t-1}$: upper values for $K = 1$ | | | |
| | Constant | | Constant and trend | | Constant | Constant and trend | Constant | | Constant and trend | |
| | Lower value | Upper value | Lower value | Upper value | | | Lower value | Upper value | Lower value | Upper value |
| 10% | 3.02 | 3.51 | 4.05 | 4.49 | 2.90 | 3.41 | −2.57 | −2.91 | −3.13 | −3.40 |
| 5% | 3.62 | 4.16 | 4.68 | 5.15 | 3.23 | 3.71 | −2.86 | −3.22 | −3.41 | −3.69 |
| 1% | 4.94 | 5.58 | 6.10 | 6.73 | 3.82 | 4.30 | −3.43 | −3.82 | −3.96 | −4.26 |

*, **, *** indicates the 1, 5, and 10 percent significance levels, respectively. $F_{PSS}$ tests the null hypothesis of no cointegration ($\alpha_2 = \alpha_3 = 0$) against the alternative of cointegration. $t_{BDM}$ tests null hypothesis of no-cointegration ($H_0 : \alpha_1 = 0$) against the alternative of cointegration. *LM* is Breusch-Godfrey serial correlation test. It has a Chi-square distribution ($\chi^2$) with 12 degrees of freedom. The 10%, 5%, and 1% critical values are 18.55, 21.03, and 26.22. *RESETtest* is Ramsey's test for functional misspecification. It is distributed as $\chi^2$ with 1 degree of freedom. The 10%, 5%, and 1% critical values are 2.71, 3.84, and 6.63. *p-values* are in square bracket. Numbers in square brackets represent probabilities. Numbers in parentheses are absolute *t-values*. *CUSUM test* is the cumulative sum of recursive residuals and *CUSUMQ test* is the cumulative sum of squares of recursive residuals. The tests are used to determine the stability of the short-term and long-term coefficients. N.S indicates an insignificant variable at conventional significance levels and hence, not included in the model.
Source: for $F_{PSS}$ and *t-value* for significance of $ECM_{t-1}$: Pesaran et al. (2001); for $t_{BDM}$: Banerjee et al. (1998). Shin et al. (2013) argue that due to the dependence structure between the partial sums $e_t^{x+}$ and $ex_t^-$, the exact value of the number of lags ($k$) is not clear. However, they argue that $ex_t^+$ and $ex_t^-$ should be treated as one variable.

insignificant long-run relationship are consistent with previous studies (see, for example, Morley and Pentecost, 2000; Nieh and Lee, 2001; Bahmani-Oskooee and Saha, 2016, 2018). Lastly, the diagnostic checks in Panel C show no evidence of either residual serial correlation or functional misspecification, and the estimated coefficients are generally stable.

We now address whether the lack of cointegration and an insignificant long-run relationship between stock prices and exchange rates are due to neglected nonlinearities. To this end, we apply the NARDL model in equation (6) and present the results in Table 4. Panel A provides the short-run estimates, Panel B provides the long-run estimates, and Panel C gives some diagnostic checks. The short-run effects of exchange rate increases ($\Delta ex^{+}$) and decreases ($\Delta ex^{-}$) on stock prices are given by the sign and significance of $\gamma_{2j}^{+}$ and $\gamma_{3j}^{-}$, respectively. The increases (decreases) imply appreciation (depreciation) in the currency of the G7 country being studied. The results show that while currency appreciations have short-run significant effects on the stock prices of all the G7 countries, except Italy, depreciations have short-run significant effects on the stock prices of only Canada. $\gamma_{2j}^{+}$ is statistically negative in France, Germany, Japan, the US, and the UK, indicating that currency appreciations ($\Delta ex^{+}$) in these countries lead to lower stock prices. The results for Canada show that currency appreciations lead to higher stock prices and depreciations lead to lower stock prices. Thus, the results suggest that short-run currency appreciations and depreciations have an asymmetric effect on the stock prices of all the G7 countries, except in Italy. For instance, in France, Germany, Japan, the US, and the UK, currency appreciations have short-run significant effects, but currency depreciations do not affect stock prices. Short-run asymmetry is also observed in Canada because $\Delta ex_t^{+}$ and $\Delta ex_t^{-}$ take different lags, and $\gamma_{2j}^{+} \neq \gamma_{3j}^{-}$.

An interesting finding for Japan is the alternating sign of the short-run coefficient estimates of yen appreciation ($\Delta ex_t^{+}$) from negative to positive. This could occur if expectations of market participants change frequently and thereby generate changes in sign.[6] What could be causing this change in expectations? Japan had experienced a prolonged deflation and liquidity trap. Stimulating aggregate demand in such an economy requires a drastic change in market expectations (Kondo et al., 2020). To this end, the Prime Minister of Japan Shinzo Abe implemented in December 2012 a set of policy measures, known as "Abenomics" (Fukuda, 2015). Kondo et al. (2020) and Fukuda (2015) examine the effects of various news shocks about Abenomics and the role of changes in investors' expectations on Japan's stock market and the yen exchange rate and find evidence of heterogeneity between domestic and foreign investors' expectations. For example, using intra-daily data, Fukuda (2015) finds that different news shocks about Abenomics had more significant impacts on stock prices and exchange rates in Japan in the nighttime than in the daytime. Their results show that although foreign investors were aggressively buying Japanese stocks and selling yen, local investors were not participating as much. Kondo et al. (2020) investigate whether Abenomics was successful in causing abrupt changes in expectations in the Japanese stock markets. Using daily data over the period January 2010 – December 2018, they find that foreign investors reacted more strongly to the launch of Abenomics than local investors and they were more aggressive in purchasing Japanese stocks than Japanese investors. This difference in the reactions between foreign and Japanese investors reflects differing expectations and views about the Japanese economy. While foreign investors seem to have more optimistic views about the health and recovery of the Japanese economy, Japanese investors seem to be more pessimistic (Fukuda, 2015). Thus, the various news shocks about the launch and implementation of Abenomics and the asymmetry in expectations between foreign investors and Japanese investors may have been responsible for this change in market participants' expectations.[7]

For these short-run asymmetric effects to last into the long-run, the variables must be cointegrated and the long-run coefficient estimates associated with currency appreciations ($ex_t^{+}$) and depreciations ($ex_t^{-}$) must be statistically significant. Combining the results from $F_{PSS}$, $t_{BDM}$, and $ECM_{t-1}$ in Panel C, we can gather that cointegration is supported for Canada, Germany, the US, and the UK. However, none of the long-run coefficient estimates associated with $ex_t^{+}$ and $ex_t^{-}$ is statistically significant. In terms of the diagnostic checks, the models pass the tests of serial correlation and functional misspecification, and the estimated coefficients have been stable. Overall, the results from testing the flow-oriented approach using linear and nonlinear ARDL models show that the effects of exchange rate changes on the G7 countries' stock prices are mainly short-run, with an insignificant long-run relationship and little evidence of cointegration.

### 4.4. *The stock-oriented model*

Given the results from testing flow-oriented model in the previous section and the lack of long-run significant relationship between stock prices and exchange rates in all the cases, we examine in this section the stock-oriented model (or the portfolio balance approach). We do this by estimating the linear and nonlinear ARDL models in equations (7) and (8). Tables 5 and 6 present the results from the linear and nonlinear ARDL models. Panel A gives short-run estimates, Panel B gives the long-run estimates, and Panel C provides some diagnostic checks.

Panel A in Table 5 shows that changes in stock prices have short-run significant effects on the exchange rates of Canada, France, Japan, the US., and the UK. These short-run effects last into the long-run in only Germany and the UK since the long-run coefficient estimates for these countries are significant at the 10% significance level. However, examining $F_{PSS}$, $t_{BDM}$ and $ECM_{t-1}$ shows that

[6] We are grateful to Mohsen Bahmani-Oskooee for his valuable comment and discussion on this point.

[7] This change in expectations and the resulting change in sign of short-run coefficient estimates from 2012 onward appear to have been caused by the implementation of Abenomics in late 2012. An estimation of the NARDL model for Japan for the period 1973:01–2011:12 (excluding the period when Abenomics was implemented) showed no evidence of changes in sign for short-run estimates. All unreported results are available upon request from the authors.

**Table 4**
Nonlinear ARDL model: short and long-run coefficient estimates (flow-oriented model).

| A. Short-run estimates | | | | | | | |
|---|---|---|---|---|---|---|---|
| | Canada | France | Germany | Italy | Japan | UK | US |
| Constant | 0.380(3.925)* | 0.169(1.808)*** | 0.207(1.969)** | 0.526(2.628)* | 0.108(2.113)** | 0.195(2.469)** | 0.116(3.548)* |
| *Trend* | 0.001(2.544)** | N.S | N.S | N.S | N.S | N.S | N.S |
| $DV_{85}$ | N.S | N.S | N.S | N.S | 0.036(2.080)** | N.S | N.S |
| $DV_{97}$ | N.S | 0.022(1.702)*** | N.S | N.S | N.S | N.S | 0.028(2.665)* |
| $DV_{03}$ | N.S | N.S | 0.039(1.953)*** | N.S | N.S | N.S | N.S |
| $DV_{07}$ | N.S | −0.021(1.813)*** | −0.023(1.923)*** | N.S | −0.022(2.121)** | 0.020(1.819)*** | −0.020(2.418)** |
| $DV_{13}$ | N.S | N.S | N.S | N.S | N.S | N.S | N.S |
| $p_{t-1}$ | −0.055(3.807)* | −0.021(1.764)*** | −0.025(1.721)*** | −0.050(2.649)* | −0.011(1.995)** | −0.025(2.389)** | −0.024(3.227)* |
| $ex^{+}_{t-1}$ | −0.013(0.720) | 0.126(1.046) | 0.233(2.467)** | −0.116(0.767) | 0.003(0.258) | −0.015(0.482) | 0.007(0.348) |
| $ex^{-}_{t-1}$ | 0.031(0.948) | 0.111(0.833) | 0.205(1.894)*** | −0.026(0.169) | 0.002(0.147) | −0.025(0.990) | −0.019(1.086) |
| $\Delta p_{t-1}$ | 0.054(1.134) | | | 0.071(1.091) | | | |
| $\Delta p_{t-2}$ | | | | −0.046(0.728) | | | |
| $\Delta p_{t-3}$ | | | | 0.127(1.990)** | | | |
| $\Delta p_{t-4}$ | | | | 0.127(2.000)** | | | |
| $\Delta ex^{+}_{t}$ | 0.500(1.882)*** | −0.840(0.994) | −1.146(1.598) | | −0.586(3.789)* | −0.441(1.736)*** | −0.964(4.665)* |
| $\Delta ex^{+}_{t-1}$ | | −1.544(1.812)*** | −0.265(0.361) | | 0.307(1.898)*** | 0.381(1.509) | |
| $\Delta ex^{+}_{t-2}$ | | 0.228(0.268) | 0.086(0.117) | | −0.331(2.032)** | | |
| $\Delta ex^{+}_{t-3}$ | | −1.518(1.779)*** | −0.791(1.088) | | 0.117(0.715) | | |
| $\Delta ex^{+}_{t-4}$ | | 1.701(2.025)** | −0.079(0.109) | | −0.063(0.386) | | |
| $\Delta ex^{+}_{t-5}$ | | | −0.125(0.172) | | 0.435(2.818)* | | |
| $\Delta ex^{+}_{t-6}$ | | | −0.274(0.378) | | | | |
| $\Delta ex^{+}_{t-7}$ | | | −0.644(0.884) | | | | |
| $\Delta ex^{+}_{t-8}$ | | | −1.637(2.250)** | | | | |
| $\Delta ex^{+}_{t-9}$ | | | −2.011(2.800)* | | | | |
| $\Delta ex^{-}_{t}$ | 0.959(3.734)* | | | | | | |
| $\Delta ex^{-}_{t-1}$ | 0.318(1.361) | | | | | | |
| B. Long-run estimates | | | | | | | |
| *constant* | 6.898(25.296) | 7.958(17.767)* | 8.445(10.954)* | 10.552(56.384)* | 9.456(16.372)* | 7.787(34.151)* | 4.908(21.681)* |
| *Trend* | 0.009(3.871)* | N.S | N.S | N.S | N.S | N.S | N.S |
| $ex^{+}_{t}$ | −0.234(0.713) | 5.926(0.795) | 9.504(1.306) | −2.324(0.725) | 0.289(0.276) | −0.604(0.434) | 0.295(0.359) |
| $ex^{-}_{t}$ | 0.558(1.028) | 5.221(0.653) | 8.344(1.060) | −0.522(0.168) | 0.199(0.152) | −1.003(0.836) | −0.813(1.058) |
| C. Diagnostic checks | | | | | | | |
| $F_{PSS}$ | 3.78 | 2.02 | 4.96** | 1.81 | 1.63 | 3.27 | 10.00* |
| $t_{BDM}$ | −0.055(3.807)** | −0.021(1.764) | −0.025(1.721) | −0.050(2.649) | −0.011(1.995) | −0.025(2.389) | −0.024(3.227)** |
| $ECM_{t-l}$ | −0.055(3.898)** | −0.021(2.854) | −0.025(4.472)* | −0.050(2.710) | −0.011(2.557) | −0.025(3.629)** | −0.024(6.343)* |
| $\sum_{j=0}^{r}\Pi^{+}_{j}$ | 0.500[0.0604]*** | −1.973[0.2356] | −6.886[0.0018]* | ——— | −0.118[0.6993] | −0.060[0.8559] | −0.964[0.0000]* |
| $\sum_{j=0}^{s}\Pi^{-}_{j}$ | 1.277[0.0001]* | ——— | ——— | ——— | ——— | ——— | ——— |
| $W_{LR}$ | 2.534[0.1121] | 0.382[0.5370] | 0.728[0.3941] | 3.330[0.0693]*** | 0.097[0.7553] | 1.148[0.2845] | 9.754[0.0019]* |
| $W_{SR}$ | 2.519[0.1132] | 1.412[0.2356] | 9.872[0.0018]* | ——— | 0.149[0.6993] | 0.033[0.8559] | 4.665[0.0000]* |
| *LMtest* | 0.068[0.7941] | 0.278[0.5981] | 0.006[0.9407] | 0.118[0.7317] | 0.048[0.8261] | 0.132[0.7170] | 0.006[0.9359] |
| *RESETtest* | 0.106[0.7454] | 0.313[0.5765] | 0.039[0.8446] | 1.020[0.3135] | 1.279[0.2586] | 0.939[0.3330] | 4.356[0.0373]** |
| $\overline{R}^{2}$ | 0.092 | 0.044 | 0.058 | 0.044 | 0.054 | 0.031 | 0.062 |
| *CUSUM* | Unstable | Stable | Stable | Stable | Stable | Stable | Stable |
| *CUSUMQ* | Unstable | Unstable | Unstable | Unstable | Unstable | Unstable | Unstable |

*, **, *** indicates the 1, 5, and 10 percent significance levels, respectively. $W_{LR}$ is a Wald test of long-run symmetry$\delta^{+} = \delta^{-}$, where $\delta^{+} = -\eta^{+}/\rho_0$ and$\delta^{-} = -\eta^{-}/\rho_0$. $W_{SR}$ is a Wald test of short-run symmetry ($\sum_{j=0}^{r}\Pi^{+}_{j} = \sum_{j=0}^{s}\Pi^{-}_{j}$).

cointegration is supported for only Germany by $ECM_{t-1}$ and at only the 10 percent significance level. Thus, once again we find that the effect of stock prices on exchange to be mainly short-term with no evidence of cointegration or significant long-run relationship. In terms of the diagnostic checks, the results suggest that the models are free of serial correlation and functional misspecification, and the estimates are stable.

Now we explore the possibility of asymmetries in the effect of stock prices on exchange rates. The results from estimating the nonlinear ARDL model in equation (8) are given in Table 6. Panel A shows that rising ($\Delta p^{+}_{st}$) and falling ($\Delta p^{-}_{st}$) stock prices have short-run significant effects on the exchange rates of all the countries, except in Germany. These short-run effects are asymmetric because the

**Table 5**
Linear ARDL model coefficient estimates and diagnostic checks (stock-oriented model).

| A. Short-run estimates | | | | | | | |
|---|---|---|---|---|---|---|---|
| | Canada | France | Germany | Italy | Japan | UK | US |
| Constant | 0.016(0.608) | 0.112(2.523)** | −0.160(3.329)* | 0.179(2.648)* | 0.035(1.739)*** | 0.083(1.872)*** | 0.038(1.278) |
| *Trend* | N.S | N.S | 0.000(1.846)*** | N.S | 0.000(1.720)*** | −0.000(2.303)** | N.S |
| $DV_{85}$ | N.S | N.S | N.S | N.S | 0.024(3.455)* | N.S | −0.010(2.304)** |
| $DV_{97}$ | N.S | N.S | N.S | N.S | N.S | N.S | N.S |
| $DV_{03}$ | N.S | N.S | N.S | N.S | N.S | N.S | −0.008(1.699)*** |
| $DV_{07}$ | N.S | N.S | 0.002(1.717)*** | 0.002(1.984)** | N.S | −0.006(1.975)** | N.S |
| $DV_{13}$ | −0.007(1.927)*** | N.S | N.S | N.S | N.S | N.S | N.S |
| $ex_{t-1}$ | −0.005(0.904) | −0.023(2.389)** | −0.031(3.109)* | −0.038(2.695)** | −0.014(2.500)** | −0.026(2.841)* | −0.007(1.275) |
| $p_{t-1}$ | 0.001(0.812) | −0.001(0.776) | −0.003(2.033)** | −0.001(0.613) | 0.002(0.841) | 0.007(1.984)** | −0.000(0.671) |
| $\Delta ex_{t-1}$ | 0.168(3.681)* | 0.238(4.528)* | 0.337(6.658)* | 0.295(4.796)* | 0.308(7.231)* | 0.230(4.932)* | 0.347(8.211)* |
| $\Delta ex_{t-2}$ | | −0.116(2.162)** | −0.094(1.849)*** | | −0.049(1.100) | | −0.082(1.923)*** |
| $\Delta ex_{t-3}$ | | 0.122(2.338)** | | | 0.063(1.430) | | |
| $\Delta ex_{t-4}$ | | | | | −0.033(0.758) | | |
| $\Delta ex_{t-5}$ | | | | | −0.015(0.338) | | |
| $\Delta ex_{t-6}$ | | | | | −0.073(1.662)*** | | |
| $\Delta ex_{t-7}$ | | | | | 0.024(0.541) | | |
| $\Delta ex_{t-8}$ | | | | | 0.062(1.418) | | |
| $\Delta ex_{t-9}$ | | | | | 0.075(1.709)*** | | |
| $\Delta ex_{t-10}$ | | | | | −0.050(1.130) | | |
| $\Delta ex_{t-11}$ | | | | | 0.121(2.893)* | | |
| $\Delta p_t$ | 0.067(4.925)* | −0.009(1.645) | | | −0.065(3.835)* | −0.033(1.924)*** | −0.051(3.714)* |
| $\Delta p_{t-1}$ | 0.083(5.979)* | −0.012(0.163)** | | | −0.036(2.080)** | 0.002(0.135) | 0.003(0.206) |
| $\Delta p_{t-2}$ | −0.017(1.231) | | | | | 0.051(2.977)* | 0.028(2.019)** |
| $\Delta p_{t-3}$ | −0.010(0.771) | | | | | | |
| $\Delta p_{t-4}$ | −0.006(0.424) | | | | | | |
| $\Delta p_{t-5}$ | 0.017(1.305) | | | | | | |
| $\Delta p_{t-6}$ | −0.021(1.589) | | | | | | |
| $\Delta p_{t-7}$ | −0.014(1.088) | | | | | | |
| $\Delta p_{t-8}$ | −0.033(2.525)** | | | | | | |
| B. Long-run estimates | | | | | | | |
| *constant* | 3.168(0.501) | 4.791(16.754)* | 5.130(17.049)* | 4.734(19.500)* | 2.528(1.895)*** | 3.147(3.447)* | 5.216(10.981)* |
| *Trend* | N.S | N.S | 0.001(1.826)*** | N.S | 0.001(2.965)* | −0.001(2.822)* | N.S |
| $p_t$ | 0.139(0.589) | −0.025(0.710) | −0.088(1.886)*** | −0.015(0.638) | 0.139(0.956) | 0.251(1.963)*** | −0.060(0.802) |
| C. Diagnostic checks | | | | | | | |
| $F_{PSS}$ | 0.59 | 2.44 | 4.14 | 2.44 | 2.47 | 3.20 | 0.71 |
| $t_{BDM}$ | −0.005(0.904) | −0.023(2.389) | −0.031(3.109) | −0.038(2.695) | −0.014(2.500) | −0.026(2.841) | −0.007(1.275) |
| $ECM_{t-l}$ | −0.005(1.336) | −0.023(2.712) | −0.031(3.535)*** | −0.038(2.714) | −0.014(2.726) | −0.026(3.107) | −0.007(1.470) |
| *LMtest* | 1.933[0.1651] | 0.010[0.9207] | 1.320[0.2513] | 0.1249[0.2649] | 1.453[0.2286] | 0.637[0.4254] | 0.001[0.9707] |
| *RESETtest* | 0.475[0.4912] | 1.644[0.2006] | 2.784[0.0960]*** | 2.730[0.0998]*** | 1.701[0.1928] | 0.761[0.3835] | 0.281[0.5965 |
| $\overline{R}^2$ | 0.192 | 0.097 | 0.124 | 0.103 | 0.180 | 0.108 | 0.156 |
| *CUSUM* | Stable | Stable | Stable | Stable | Stable | Stable | Stable |
| *CUSUMQ* | Stable | Unstable | Unstable | Stable | Stable | Stable | Unstable |

*, **, *** indicates the 1, 5, and 10 percent significance levels, respectively.

variables $\Delta p^+_{st}$ and $\Delta p^-_{st}$ take different lag orders across countries and the sign or size of short-run estimates associated with the same lag varies by country. For example, while falling stock prices have short-run significant effects on the exchange rates of Italy and Japan, rising prices are insignificant. In France and the UK, rising stock prices have short-run significant effect, but falling prices have no effect on exchange rates. Furthermore, the null hypothesis of short-run impact symmetry is rejected for France, Italy, and Japan.

Short-run effects last into the long-run in France, Germany, Italy, and the UK. Rising $(p^+_t)$ and falling $(p^-_t)$ stock prices have a long-run significant effect on the exchange rates in these countries. These long-run effects are asymmetric in Germany, Italy and the UK since the null hypothesis of long-run symmetry is rejected for these countries. To make sure that these long-run results are not spurious, we examine cointegration between the variables. The results of cointegration tests in Panel C of Table 6 show that cointegration is supported by at least one of the tests ($F_{PSS}$, $t_{BDM}$ or $ECM_{t-1}$) in each of the four countries. The diagnostic checks suggest that the models are free of serial correlation and functional misspecification. Also, the coefficient estimates are relatively stable.

**Table 6**
Nonlinear ARDL model: short and long-run coefficient estimates (stock-oriented model).

| | Canada | France | Germany | Italy | Japan | UK | US |
|---|---|---|---|---|---|---|---|
| A. Short-run estimates | | | | | | | |
| Constant | 0.020(0.798) | 0.203(3.629) | 0.156(3.371)* | 0.228(3.758)* | 0.061(3.086)* | 0.135(2.920)* | 0.045(1.483) |
| *Trend* | N.S | N.S | N.S | 0.004(3.949)* | N.S | N.S | N.S |
| $DV_{85}$ | N.S | N.S | N.S | N.S | 0.025(3.649)* | N.S | −0.010(2.260)** |
| $DV_{97}$ | N.S | 0.004(3.035)* | 0.003(1.676)*** | N.S | N.S | N.S | N.S |
| $DV_{03}$ | N.S | 0.003(1.989)** | N.S | 0.004(2.560)** | N.S | N.S | −0.008(1.673)*** |
| $DV_{07}$ | N.S | 0.003(2.316)** | 0.002(1.787)*** | N.S | N.S | −0.005(2.001)** | N.S |
| $DV_{13}$ | −0.007(1.831)** | N.S | N.S | 0.004(2.524)** | N.S | N.S | N.S |
| $ex_{t-1}$ | −0.005(0.859) | −0.044(3.571)* | −0.034(3.344)* | −0.073(4.272)* | −0.017(3.078)* | −0.028(2.992)* | −0.009(1.429) |
| $p^{+}_{t-1}$ | 0.003(0.756) | −0.002(1.968)** | −0.003(2.579)** | −0.014(3.452)* | 0.003(1.203) | 0.006(2.065)** | 0.001(0.434) |
| $p^{-}_{t-1}$ | 0.003(0.705) | −0.002(1.940)*** | −0.004(2.586)** | 0.003(1.715)*** | 0.002(0.819) | 0.009(2.254)** | 0.002(0.555) |
| $\Delta ex_{t-1}$ | 0.163(3.582)* | 0.205(3.872)* | 0.333(6.587)* | 0.280(4.582)* | 0.296(6.974)* | 0.232(4.956)* | 0.346(8.214)* |
| $\Delta ex_{t-2}$ | | −0.119(2.226)** | −0.098(1.929)*** | | −0.055(1.247) | | −0.084(1.969)** |
| $\Delta ex_{t-3}$ | | 0.106(2.022)** | | | 0.065(1.490) | | |
| $\Delta ex_{t-4}$ | | | | | −0.034(0.771) | | |
| $\Delta ex_{t-5}$ | | | | | −0.021(0.493) | | |
| $\Delta ex_{t-6}$ | | | | | −0.074(1.703)*** | | |
| $\Delta ex_{t-7}$ | | | | | 0.022(0.504) | | |
| $\Delta ex_{t-8}$ | | | | | 0.066(1.509) | | |
| $\Delta ex_{t-9}$ | | | | | 0.071(1.632) | | |
| $\Delta ex_{t-10}$ | | | | | 0.132(3.160)* | | |
| $\Delta ex_{t-11}$ | | | | | | | |
| $\Delta p^{+}_{t}$ | 0.067(2.411)** | −0.036(3.067)* | | | | −0.076(2.368)** | −0.048(1806)*** |
| $\Delta p^{+}_{t-1}$ | 0.053(1.880)** | −0.034(3.358)* | | | | 0.006(0.194) | |
| $\Delta p^{+}_{t-2}$ | −0.076(2.706)* | −0.008(0.789) | | | | 0.067(2.097)** | |
| $\Delta p^{+}_{t-3}$ | | 0.000(0.041) | | | | | |
| $\Delta p^{+}_{t-4}$ | | −0.004(0.390) | | | | | |
| $\Delta p^{+}_{t-5}$ | | −0.029(2.920)* | | | | | |
| $\Delta p^{-}_{t}$ | 0.076(3.237)* | 0.013(1.294) | | 0.018(2.243)** | −0.104(3.929)* | | −0.050(2.030)** |
| $\Delta p^{-}_{t-1}$ | 0.098(4.443)* | | | 0.008(0.974) | −0.051(1.893)*** | | 0.009(0.383) |
| $\Delta p^{-}_{t-2}$ | 0.017(0.734) | | | | −0.050(1.837)*** | | 0.042(1.855)*** |
| $\Delta p^{-}_{t-3}$ | −0.028(1.315) | | | | | | |
| $\Delta p^{-}_{t-4}$ | −0.009(0.412) | | | | | | |
| $\Delta p^{-}_{t-5}$ | 0.031(1.504) | | | | | | |
| $\Delta p^{-}_{t-6}$ | −0.051(2.495)** | | | | | | |
| $\Delta p^{-}_{t-7}$ | −0.028(1.361) | | | | | | |
| $\Delta p^{-}_{t-8}$ | −0.060(2.940)* | | | | | | |
| B. Long-run estimates | | | | | | | |
| *constant* | 4.173(7.795)* | 4.637(145.985)* | 4.606(168.77)* | 3.113(12.401)* | 3.509(26.595)* | 4.755(46.136)* | 5.124(22.177)* |
| *Trend* | N.S | N.S | N.S | 0.005(5.867)* | N.S | N.S | N.S |
| $p^{+}_{t}$ | 0.568(0.531) | −0.047(1.990)** | −0.092(2.293)** | −0.192(4.883)* | 0.169(1.503) | 0.224(2017)** | 0.089(0.430) |
| $p^{-}_{t}$ | 0.716(0.511) | −0.054(2.015)** | −0.119(2.334)** | 0.037(1.723)*** | 0.108(0.917) | 0.320(2.308)** | 0.172(0.555) |
| C. Diagnostic checks | | | | | | | |
| $F_{PSS}$ | 0.54 | 5.99* | 3.79*** | 4.92*** | 4.10*** | 2.66 | 1.02 |
| $t_{BDM}$ | −0.005(0.859) | −0.044(3.571)** | −0.034(3.344)** | −0.014(3.452)* | −0.017(3.078)*** | −0.028(2.992)*** | −0.009(1.429) |
| $ECM_{t-l}$ | −0.005(1.475) | −0.044(4.918)* | −0.034(3.910)* | −0.073(4.465)* | −0.017(4.059)* | −0.028(3.273)** | −0.009(2.030) |
| $\sum_{j=0}^{r}\Pi_j^{+}$ | 0.044[0.3754] | −0.111[0.0002]* | ——— | ——— | ——— | −0.003[0.9693] | −0.048[0.0714]*** |
| $\sum_{j=0}^{s}\Pi_j^{-}$ | 0.046[0.3541] | 0.013[0.1967] | ——— | 0.026[0.0245]** | −0.205[0.0000]* | ——— | 0.001[0.9696] |
| $W_{LR}$ | 0.300[0.5842] | 1.738[0.1883] | 5.9939[0.0153]** | 15.289[0.0001]* | 4.414[0.0361]** | 6.216[0.0130]** | 0.571[0.4503] |
| $W_{SR}$ | 0.002[0.9636] | 12.852[0.0004]* | ——— | 2.264[0.0245]** | 24.180[0.0000]* | 0.002[0.9693] | 1.166[0.2807] |
| *LMtest* | 1.041[0.3082] | 0.110[0.7405] | 0.912[0.3403] | 0.852[0.3569] | 0.050[0.8223] | 0.756[0.3852] | 0.064[0.7998] |
| *RESETtest* | 0.259[0.6110] | 0.842[0.3536] | 1.580[0.2096] | 0.730[0.3938] | 2.809[0.09943]*** | 0.453[0.5015] | 0.694[0.4053] |
| $\overline{R}^2$ | 0.208 | 0.139 | 0.131 | 0.143 | 0.190 | 0.105 | 0.155 |
| *CUSUM* | Stable | Stable | Stable | Stable | Stable | Stable | Stable |
| *CUSUMQ* | Stable | Unstable | Unstable | Unstable | Stable | Stable | Unstable |

*, **, *** indicates the 1, 5, and 10 percent significance levels, respectively.

Allowing for asymmetries in the effects of stock prices on the exchange rates, we find evidence of long-run significant asymmetric relationship between stock prices and exchange rates that is not spurious in France, Germany, Italy, and the UK. If instead, we had relied on our findings from the linear ARDL model, we would have concluded that stock prices have no long-run significant effect on the exchange rates of these four countries.

The results from both the linear and nonlinear ARDL models in Tables 3 – 6 are summarized in Table 7. We find short-run evidence supporting both the flow-oriented and portfolio balance approaches. There is no long-run evidence to confirm the flow-oriented approach that changes in exchange rates cause changes in stock prices, based upon either the linear or the nonlinear ARDL model. Similarly, there is no support for the portfolio balance approach using the linear ARDL model. In contrast, our results from the nonlinear ARDL model provide long-run support for the portfolio balance approach in France, Germany, Italy, and the UK. In these four countries, rising ($p^+$) and falling ($p^-$) stock prices have significant and nonspurious long-run effects on exchange rates, but not always in the direction predicted by the portfolio balance approach.

Changes in stock prices have both short-run and long-run effects on the exchange rates of France, Italy, and the UK, but have only short-run effects on the exchange rates of Canada, Japan, and the US, and only a long-run effect on the exchange rate of Germany. The results for Canada that rising (falling) stock prices lead to short-run dollar appreciation (depreciation) is consistent with the portfolio balance approach. Similarly, the finding for the UK that rising (falling) stock prices lead to long-run pound appreciation (depreciation) is consistent with the portfolio approach. For Italy, falling stock prices lead to currency depreciation in both the short-run and long-run, which is consistent with the portfolio balance approach. However, rising stock prices lead to currency depreciation in the long-run, which is inconsistent with this approach.[8] The results for France and Germany that rising (falling) stock prices lead to currency depreciation (appreciation) in the long-run, is not consistent with the portfolio approach. In Japan, falling stock prices lead to short-run yen appreciation, which is again inconsistent with the portfolio balance approach. Finally, rising (falling) US stock prices lead to dollar depreciation (appreciation)) the dollar in the short-run, which is not consistent with the portfolio balance approach.

Our findings for France, Germany, and the US, that rising (falling) stock prices lead to currency depreciation (appreciation) violate the portfolio balance approach but are in line with results obtained by Ajayi and Mougoue (1996) who examined the G7 and the Netherlands. The authors found that rising stock prices caused short-run currency depreciation in all cases. They argue that a booming domestic stock market indicates that the economy is expanding, which can increase inflation expectations. Higher expected inflation negatively affects international investors who will reduce their investments and their demand for the domestic currency. Consequently, higher inflation expectation puts downward pressure on the domestic currency, and it will depreciate. Thus, rising stock price can lead to currency depreciation through inflationary expectations.

Although our short-run findings of a bi-directional relationship between stock prices and exchange rates are consistent with much of the literature, our long-run results show a more complicated relationship regarding the impact of stock prices on exchange rates than found in previous studies. For example, our results contradict those of Morley and Pentecost (2000), Nieh and Lee (2001), Yang and Doong (2004), and Bahmani-Oskooee and Saha (2018) who find no evidence of cointegration for the G7 countries. Differences are probably due to our use of nonlinear ARDL models, allowing for structural breaks, and because we use more recent data. For example, Salisu and Ndako (2018) point out that the validity of the portfolio balance approach is more evident after the Global Financial Crisis than in earlier periods. On the other hand, our results complement those of Kollias et al. (2016) who find evidence of evidence of cointegration in France, Germany, Italy, and the UK based upon a threshold cointegration methodology. Finally, our results for the US of a short-run significant relationship between stock prices and exchange rates, but lack of a long-run significant relationship are in line with previous research by Bahmani-Oskooee and Sohrabian (1992), Ratner (1993), and Chow et al. (1997).

### 4.5. *Robustness of the results*

To check the robustness of our previous results, we now utilize a simple VAR model to examine the interaction between the exchange rate and stock prices in each country. The VAR model is used to test for cointegration between variables using the Johansen (1998) procedure and then the direction of causality between the variables is determined by the Granger (1969) causality test. To this end, a VAR model is estimated with the optimal number of lags selected by the smallest AIC. Four different VAR models are estimated: 1) a linear VAR model containing the variables $ex$ and$p$; 2) a nonlinear VAR model containing the variables$p$, $ex^+$ and$ex^-$; 3) a nonlinear VAR model containing the variables$ex$, $p^+$ and$p^-$; and 4) a nonlinear VAR containing the variables$ex^+$,$ex^-$, $p^+$ and$p^-$. Tables 8 – 11 report the results from Johansen cointegration and Granger causality tests, and Table 12 provides a summary of these tables.

Starting with Table 8, the results show no evidence of cointegration between exchange rates ($ex$) and stock prices($p$), which is in line with the results from the linear ARDL models. A Granger causality test shows evidence of causality from $p$ to $ex$ in Canada, France, Japan, and the UK, with no evidence of reverse causality. This provides support for the portfolio balance model that stock prices affect exchange rates. We find evidence of Granger causality from $ex$ to$p$, supporting the flow-oriented model for Italy, but the relationship is not supported in the other six G7 countries. Allowing for nonlinearity by separating currency appreciations ($ex^+$) from depreciations ($ex^-$) and examining cointegration among$p$, $ex^+$ and$ex^-$, the results in Table 9 show no evidence of cointegration at the 5 percent significance level in any G7 country. However, at the 10 percent significance level, evidence of cointegration emerges for Canada. Granger causality tests indicate unidirectional causality from $p$ to either $ex^+$ or$ex^-$, or both in Canada, Germany, Japan, the UK, and the US. This provides support for the portfolio balance model in these countries. We next introduce nonlinearities into stock prices by

[8] Italy may behave differently than France or Germany because its economy is a smaller portion of the Euro zone. Stock price changes in the Italian market may therefore have little impact on the euro relative to stock price changes in France or Germany.

**Table 7**
Summary of the linear and nonlinear ARDL models (Tables 3 – 6).

| | Flow-oriented model | | | | | | | |
|---|---|---|---|---|---|---|---|---|
| | Linear ARDL Model: $p = f(ex)$ | | | Nonlinear ARDL Model: $p = f(ex^{+}, ex^{-})$ | | | | |
| Country | Short-run effect of $ex$ | Long-run effect of $ex$ | Cointegration | Short-run effect of $ex^{+}$ | Short-run effect of $ex^{-}$ | Long-run effect of $ex^{+}$ | Long-run effect of $ex^{-}$ | Cointegration |
| Canada | + | N.S | No | + | + | N.S | N.S | Yes |
| France | – | N.S | No | – | N.E | N.S | N.S | No |
| Germany | – | N.S | No | – | N.E | N.S | N.S | Yes |
| Italy | N.E | N.S | No | N.E | N.E | N.S | N.S | No |
| Japan | – | N.S | No | – | N.E | N.S | N.S | No |
| UK | – | N.S | Yes | – | N.E | N.S | N.S | Yes |
| US | – | N.S | No | – | N.E | N.S | N.S | Yes |
| | Stock-oriented model | | | | | | | |
| | Linear ARDL Model: $ex = f(p)$ | | | Nonlinear ARDL Model: $ex = f(p^{+}, p^{-})$ | | | | |
| Country | Short-run effect of $p$ | Long-run effect of $p$ | Cointegration | Short-run effect of $p^{+}$ | Short-run effect of $p^{-}$ | Long-run effect of $p^{+}$ | Long-run effect of $p^{-}$ | Cointegration |
| Canada | + | N.S | No | + | + | N.S | N.S | No |
| France | – | N.S | No | – | N.S | – | – | Yes |
| Germany | N.E | – | No | N.E | N.E | – | – | Yes |
| Italy | N.E | N.S | No | N.E | + | – | + | Yes |
| Japan | – | N.S | No | N.E | – | N.S | N.S | Yes |
| UK | – | + | No | – | N.E | + | + | Yes |
| US | – | N.S | No | – | – | N.S | N.S | No |

+ (−) means a positive (negative) coefficient estimate associated with the independent variable(s). N.S means statistically insignificant coefficient estimate at conventional significance levels. N.E means the independent variable has no effect on the dependent variable. For the short-run effect, the + (−) sign stands for cumulative short-run effect of the independent variable on the dependent variable. Cointegration is reported only if the null hypothesis of no-cointegration is rejected at the 5 percent significance level or lower by at least one of the tests $F_{PSS}$, $t_{BDM}$, or $ECM_{t-l}$..

**Table 8**
Johansen cointegration and Granger causality tests between $p$ and $ex$ (linear analysis).

| | Johansen cointegration test | | | | Granger causality test | | |
|---|---|---|---|---|---|---|---|
| | | | | | Null hypothesis | | |
| Country | Lags | $H_0$ | $\lambda_{Trace}$ | $\lambda_{Max}$ | Lags | $p \nRightarrow ex$ | $ex \nRightarrow p$ |
| Canada | 2 | $r = 0$ | 5.31 | 3.40 | 1 | 37.60 | 2.16 |
| | | $r \leq 1$ | 1.91 | 1.91 | | [0.0000]* | [0.1416] |
| France | 2 | $r = 0$ | 10.11 | 6.39 | 1 | 6.06 | 0.31 |
| | | $r \leq 1$ | 3.72 | 3.72 | | [0.0138]** | [0.5799] |
| Germany | 2 | $r = 0$ | 10.20 | 6.16 | 1 | 2.31 | 0.40 |
| | | $r \leq 1$ | 4.04 | 4.04 | | [0.1286] | [0.5276] |
| Italy | 2 | $r = 0$ | 5.23 | 3.17 | 1 | 0.04 | 2.83 |
| | | $r \leq 1$ | 2.06 | 2.06 | | [0.8461] | [0.0927]*** |
| Japan | 2 | $r = 0$ | 8.83 | 5.48 | 1 | 3.69 | 0.08 |
| | | $r \leq 1$ | 3.35 | 3.35 | | [0.0547]*** | [0.7718] |
| UK | 3 | $r = 0$ | 12.64 | 8.87 | 3 | 12.76 | 0.73 |
| | | $r \leq 1$ | 3.77 | 3.77 | | [0.0017]* | [0.6944] |
| US | 3 | $r = 0$ | 4.78 | 4.67 | 2 | 4.06 | 0.95 |
| | | $r \leq 1$ | 0.11 | 0.11 | | [0.1311] | [0.6213] |

*, **, *** indicates the 1, 5, and 10 percent significance levels, respectively. Lags reported are the number of lags in the VAR model that results in the smallest AIC. If there is evidence of serial autocorrelation, number of lags is increased until there is no evidence of serial autocorrelation. *p-values* are square bracketed. $p \nRightarrow ex$ means $p$ does not Granger cause $ex$, and $ex \nRightarrow p$ means $ex$ does not Granger cause $p$. For the Granger causality test, a VAR model in first differences is estimated if there is no evidence of cointegration, and an error-correction model is estimated if there is evidence of cointegration. $\lambda_{Trace}$ and $\lambda_{Max}$ are the Trace statistic and the Max-Eigen statistic for the null hypothesis of no-cointegration ($H_0$).

separating stock price increases ($p^{+}$) from the decreases ($p^{-}$), and we examine cointegration among the variables $ex$, $p^{+}$ and $p^{-}$. Results Table 10 reveal evidence of cointegration in Canada and the UK, but not in the other five countries.

Like the results from our nonlinear ARDL models, Granger causality tests mostly indicate unidirectional causality from stock prices to exchange rates. The tests show unidirectional causality from either $p^{+}$ or $p^{-}$, or both, to $ex$ in Canada, France, Japan, the UK, and the US. However, in the UK, we also find evidence of causality from $ex$ to $p^{+}$. Lastly, we allow for nonlinearities in both $ex$ and $p$ and examine the cointegration among the variables $ex^{+}$, $ex^{-}$, $p^{+}$ and $p^{-}$. Table 11 finds support for cointegration only in Canada and the UK.

**Table 9**
Johansen cointegration and Granger causality tests between $p$ and $ex^{+}$ and $ex^{-}$ (nonlinear analysis).

| Country | Johansen cointegration test | | | | Granger causality test | | | | |
|---|---|---|---|---|---|---|---|---|---|
| | Lags | $H_0$ | $\lambda_{Trace}$ | $\lambda_{Max}$ | | Null hypothesis | | | |
| | | | | | Lags | $p \nRightarrow ex^{+}$ | $p \nRightarrow ex^{-}$ | $ex^{+} \nRightarrow p$ | $ex^{-} \nRightarrow p$ |
| Canada | 2 | $r = 0$ | 28.58*** | 20.04*** | 2 | 20.30 | 37.38 | 1.01 | 3.28 |
| | | $r \leq 1$ | 8.55 | 6.36 | | [0.0000]* | [0.0000]* | [0.6033] | [0.1937] |
| | | $r \leq 2$ | 2.19 | 2.19 | | | | | |
| France | 2 | $r = 0$ | 33.60** | 23.06** | 3 | 5.42 | 4.78 | 4.38 | 2.94 |
| | | $r \leq 1$ | 10.53 | 6.10 | | [0.1434] | [0.1888] | [0.2229] | [0.4006] |
| | | $r \leq 2$ | 4.43** | 4.43** | | | | | |
| Germany | 3 | $r = 0$ | 31.10** | 14.10 | 2 | 5.73 | 2.30 | 0.46 | 0.62 |
| | | $r \leq 1$ | 16.10** | 10.03 | | [0.0571]*** | [0.3168] | [0.7946] | [0.7324] |
| | | $r \leq 2$ | 6.07** | 6.07** | | | | | |
| Italy | 2 | $r = 0$ | 15.32 | 9.25 | 1 | 0.34 | 1.11 | 2.49 | 0.04 |
| | | $r \leq 1$ | 6.08 | 4.14 | | [0.5611] | [0.2922] | [0.1143] | [0.8505] |
| | | $r \leq 2$ | 1.94 | 1.94 | | | | | |
| Japan | 2 | $r = 0$ | 19.65 | 14.42 | 1 | 6.74 | 0.10 | 0.45 | 0.18 |
| | | $r \leq 1$ | 5.22 | 3.83 | | [0.0094]* | [0.7460] | [0.5033] | [0.6706] |
| | | $r \leq 2$ | 1.40 | 1.40 | | | | | |
| UK | 2 | $r = 0$ | 22.42 | 11.69 | 3 | 7.51 | 8.72 | 5.79 | 0.52 |
| | | $r \leq 1$ | 10.73 | 9.93 | | [0.0574]*** | [0.0332]** | [0.1223] | [0.9150] |
| | | $r \leq 2$ | 0.80 | 0.80 | | | | | |
| US | 2 | $r = 0$ | 10.44 | 7.17 | 2 | 3.94 | 5.49 | 0.70 | 0.57 |
| | | $r \leq 1$ | 3.27 | 3.25 | | [0.1396] | [0.0642]*** | [0.7031] | [0.7506] |
| | | $r \leq 2$ | 0.02 | 0.02 | | | | | |

*, **, and *** indicate significance at the 1, 5, and 10 percent significance levels. *p-values* are square bracketed.$p \nRightarrow ex^{+}$ means $p$ does not Granger cause$ex^{+}$, where $ex^{+}$ and $ex^{-}$ are currency appreciation and depreciation of the currency of the G7 country.

**Table 10**
Johansen cointegration and Granger causality tests between $ex$ and $p^{+}$ and $p^{-}$ (nonlinear analysis).

| Country | Johansen cointegration test | | | | Granger causality test | | | | |
|---|---|---|---|---|---|---|---|---|---|
| | Lags | $H_0$ | $\lambda_{Trace}$ | $\lambda_{Max}$ | | Null hypothesis | | | |
| | | | | | Lags | $p^{+} \nRightarrow ex$ | $p^{-} \nRightarrow ex$ | $ex \nRightarrow p^{+}$ | $ex \nRightarrow p^{-}$ |
| Canada | 3 | $r = 0$ | 36.23* | 23.25** | 3 | 9.38 | 19.76 | 6.16 | 0.60 |
| | | $r \leq 1$ | 12.97 | 10.47 | | [0.0246]** | [0.0002]* | [0.1039] | [0.8971] |
| | | $r \leq 2$ | 2.51 | 2.51 | | | | | |
| France | 5 | $r = 0$ | 19.60 | 9.61 | 5 | 18.83 | 10.23 | 8.49 | 0.94 |
| | | $r \leq 1$ | 9.99 | 5.52 | | [0.0021]* | [0.0689]*** | [0.1312] | [0.9675] |
| | | $r \leq 2$ | 4.46** | 4.46** | | | | | |
| Germany | 3 | $r = 0$ | 20.34 | 9.87 | 2 | 0.34 | 4.40 | 0.62 | 0.17 |
| | | $r \leq 1$ | 10.47 | 8.06 | | [0.8431] | [0.1108] | [0.7331] | [0.9191] |
| | | $r \leq 2$ | 2.41 | 2.41 | | | | | |
| Italy | 3 | $r = 0$ | 20.17 | 16,32 | 2 | 0.60 | 0.65 | 3.32 | 1.78 |
| | | $r \leq 1$ | 3.85 | 3.46 | | [0.7420] | [0.7238] | [0.1901] | [0.4104] |
| | | $r \leq 2$ | 0.39 | 0.39 | | | | | |
| Japan | 3 | $r = 0$ | 29.17*** | 22.33** | 4 | 3.86 | 10.47 | 5.00 | 0.63 |
| | | $r \leq 1$ | 6.84 | 6.39 | | [0.2770] | [0.0149]** | [0.1719] | [0.8897] |
| | | $r \leq 2$ | 0.45 | 0.45 | | | | | |
| UK | 4 | $r = 0$ | 39.08* | 25.49* | 4 | 2.99 | 8.63 | 7.94 | 0.75 |
| | | $r \leq 1$ | 13.59 | 10.84 | | [0.5593] | [0.0711]*** | [0.0938]*** | [0.9449] |
| | | $r \leq 2$ | 2.75 | 2.75 | | | | | |
| US | 7 | $r = 0$ | 12.20 | 7.26 | 6 | 6.25 | 12.32 | 7.93 | 1.56 |
| | | $r \leq 1$ | 4.94 | 4.31 | | [0.3962] | [0.0552]*** | [0.2436] | [0.9552] |
| | | $r \leq 2$ | 0.63 | 0.63 | | | | | |

*, **, and *** indicate significance at the 1, 5, and 10 percent significance levels. *p-values* are square bracketed. $ex \nRightarrow p^{+}$ means $ex$ does not Granger cause$p^{+}$, where $p^{+}$ and $p^{-}$ are increases and decreases in the stock prices of the G7 country.

We find a unidirectional causality from either $p^{+}$ to either $ex^{+}$ or $ex^{-}$ (or both), or from $p^{-}$ to either $ex^{+}$ or$ex^{-}$, (or both) in Canada, France, Germany, Japan, the UK, and the US. These results reinforce and provide somewhat greater support for the portfolio balance approach than the results from nonlinear ARDL models in the previous section. In contrast, we find evidence of a unidirectional causality from $ex^{+}$ to $p^{-}$ in Italy, which provides mild support for the flow-oriented approach. Nevertheless, allowing for nonlinearities

**Table 11**
Johansen cointegration and Granger causality tests between$ex^+$,$ex^-$, $p^+$ and $p^-$ (nonlinear analysis).

| Country | Johansen cointegration test | | | | Granger causality test | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | Lags | $H_0$ | $\lambda_{Trace}$ | $\lambda_{Max.}$ | Null hypothesis | | | | | | | | |
| | | | | | Lags | $p^+ \nRightarrow ex^+$ | $p^+ \nRightarrow ex^-$ | $p^- \nRightarrow ex^-$ | $p^- \nRightarrow ex^+$ | $ex^+ \nRightarrow p^+$ | $ex^+ \nRightarrow p^-$ | $ex^- \nRightarrow p^-$ | $ex^- \nRightarrow p^+$ |
| Canada | 3 | $r = 0$ | 56.49* | 33.18* | 3 | 12.83 | 4.95 | 29.84 | 4.38 | 0.65 | 1.15 | 2.67 | 8.18 |
| | | $r \leq 1$ | 23.31 | 19.14*** | | [0.0050]* | [0.1758] | [0.0000]* | [0.2229] | [0.8850] | [0.7659] | [0.4458] | [0.0424]** |
| | | $r \leq 2$ | 4.17 | 4.17 | | | | | | | | | |
| | | $r \leq 3$ | 0.00 | 0.000 | | | | | | | | | |
| France | 4 | $r = 0$ | 40.58 | 21.44 | 3 | 4.25 | 8.91 | 2.01 | 6.43 | 3.15 | 3.90 | 4.21 | 2.61 |
| | | $r \leq 1$ | 19.14 | 10.75 | | [0.2359] | [0.0305]** | [0.5700] | [0.0927]*** | [0.3687] | [0.2726] | [0.2401] | [0.4555] |
| | | $r \leq 2$ | 8.39 | 4.76 | | | | | | | | | |
| | | $r \leq 3$ | 3.63 | 3.63 | | | | | | | | | |
| Germany | 3 | $r = 0$ | 41.55 | 16.60 | 2 | 0.26 | 0.56 | 1.76 | 5.02 | 1.62 | 1.09 | 0.42 | 3.97 |
| | | $r \leq 1$ | 24.95 | 10.96 | | [0.8778] | [0.7560] | [0.4138] | [0.0812]*** | [0.4440] | [0.5805] | [0.8096] | [0.1376] |
| | | $r \leq 2$ | 13.98*** | 8.46 | | | | | | | | | |
| | | $r \leq 3$ | 5.52** | 5.52** | | | | | | | | | |
| Italy | 2 | $r = 0$ | 64.49* | 28.82** | 4 | 0.87 | 1.33 | 1.31 | 2.49 | 4.23 | 9.40 | 8.83 | 3.79 |
| | | $r \leq 1$ | 35.67* | 21.58** | | [0.9295] | [0.8563] | [0.8595] | [0.6469] | [0.3762] | [0.0519]*** | [0.0655]*** | [0.4349] |
| | | $r \leq 2$ | 14.09*** | 11.33 | | | | | | | | | |
| | | $r \leq 3$ | 2.76*** | 2.76*** | | | | | | | | | |
| Japan | 3 | $r = 0$ | 39.85 | 23.22 | 3 | 8.64 | 2.55 | 2.00 | 15.79 | 4.22 | 0.83 | 0.81 | 0.41 |
| | | $r \leq 1$ | 16.62 | 10.16 | | [0.0343]** | [0.4658] | [0.5717] | [0.0012]* | [0.2387] | [0.8425] | [0.8476] | [0.9388] |
| | | $r \leq 2$ | 6.47 | 5.11 | | | | | | | | | |
| | | $r \leq 3$ | 1.36 | 1.36 | | | | | | | | | |
| UK | 4 | $r = 0$ | 45.94*** | 27.40*** | 3 | 5.38 | 1.23 | 9.10 | 3.92 | 3.51 | 5.54 | 1.91 | 2.63 |
| | | $r \leq 1$ | 18.54 | 10.81 | | [0.2505] | [0.8732] | [0.0587]*** | [0.4169] | [0.4760] | [0.2364] | [0.7520] | [0.6220] |
| | | $r \leq 2$ | 7.72 | 7.03 | | | | | | | | | |
| | | $r \leq 3$ | 0.69 | 0.69 | | | | | | | | | |
| US | 7 | $r = 0$ | 29.86 | 19.48 | 6 | 5.51 | 4.60 | 13.73 | 15.92 | 2.90 | 2.01 | 1.65 | 5.24 |
| | | $r \leq 1$ | 10.39 | 6.69 | | [0.4801] | [0.5961] | [0.0328]** | [0.0142]** | [0.8217] | [0.9186] | [0.9487] | [0.5136] |
| | | $r \leq 2$ | 3.70 | 3.59 | | | | | | | | | |
| | | $r \leq 3$ | 0.11 | 0.11 | | | | | | | | | |

*, **, and *** indicate significance at the 1, 5, and 10 percent significance levels. *p-values* are square bracketed. $ex \nRightarrow p^+$ means *ex* does not Granger cause$p^+$, where $p^+$ and $p^-$ are increases and decreases in the stock prices.

**Table 12**
Summary of Johansen cointegration and Granger causality tests (Tables 8 – 11).

| | Johansen cointegration tests | | | |
|---|---|---|---|---|
| Country | Cointegration between $p$ and $ex$ | Cointegration between $p$, $ex^{+}$ and $ex^{-}$ | Cointegration between $ex$, $p^{+}$ and $p^{-}$ | Cointegration between $ex^{+}$, $ex^{-}$, $p^{+}$ and $p^{-}$ |
| Canada | No | No | Yes | Yes |
| France | No | No | No | No |
| Germany | No | No | No | No |
| Italy | No | No | No | No |
| Japan | No | No | Yes | No |
| UK | No | No | Yes | Yes |
| US | No | No | No | No |

| Granger causality tests | Granger causality test between $p$ and $ex$ | | Granger causality test between $p$, $ex^{+}$ and $ex^{-}$ | | | | Granger causality test between $ex$, $p^{+}$ and $p^{-}$ | | | |
|---|---|---|---|---|---|---|---|---|---|---|
| Country | $p \Rightarrow ex$ | $ex \Rightarrow p$ | $p \Rightarrow ex^{+}$ | $p \Rightarrow ex^{-}$ | $ex^{+} \Rightarrow p$ | $ex^{-} \Rightarrow p$ | $ex \Rightarrow p^{+}$ | $ex \Rightarrow p^{-}$ | $p^{+} \Rightarrow ex$ | $p^{-} \Rightarrow ex$ |
| Canada | Yes | No | Yes | Yes | No | No | No | No | Yes | Yes |
| France | Yes | No | No | No | No | No | No | No | Yes | Yes |
| Germany | No | No | Yes | No | No | No | No | No | No | No |
| Italy | No | Yes | No | No | No | No | No | No | No | No |
| Japan | Yes | No | Yes | No | No | No | No | No | No | Yes |
| UK | Yes | No | Yes | Yes | No | No | Yes | No | No | Yes |
| US | No | No | No | Yes | No | No | No | No | No | Yes |

| | Granger causality test between $p^{+}$, $p^{-}$, $ex^{+}$ and $ex^{-}$ | | | | | | | |
|---|---|---|---|---|---|---|---|---|
| Country | $p^{+} \Rightarrow ex^{+}$ | $p^{+} \Rightarrow ex^{-}$ | $p^{-} \Rightarrow. ex^{+}$ | $p^{-} \Rightarrow. ex^{-}$ | $ex^{+} \Rightarrow p^{+}$ | $ex^{+} \Rightarrow p^{-}$ | $ex^{-} \Rightarrow p^{+}$ | $ex^{-} \Rightarrow p^{-}$ |
| Canada | Yes | No | No | Yes | No | No | Yes | No |
| France | No | Yes | Yes | No | No | No | No | No |
| Germany | No | No | Yes | No | No | No | No | No |
| Italy | No | No | No | No | No | Yes | No | Yes |
| Japan | Yes | No | Yes | No | No | No | No | No |
| UK | No | No | No | Yes | No | No | No | No |
| US | No | No | Yes | Yes | No | No | No | No |

Cointegration is reported only if the null hypothesis of no-cointegration is rejected at the 5 percent significance level or lower. $p \Rightarrow ex$ means the variable $p$ Granger causes the variable $ex$.

in the variables by separating increases from decreases, provides much stronger support for the portfolio balance approach that stock prices affect exchange rates than if linear models are used. Also, support for the portfolio balance approach based upon introducing nonlinearities into both stock prices and exchange rates is robust across techniques, based upon both ARDL models and VAR models with Granger causality tests.

## 5. Summary and policy implications

In this paper we have examined the relationship between stock prices and exchange rates in the G7 countries using monthly data up to February of 2020. Both the flow-oriented approach (that changes in exchange rates cause changes in stock prices) and the stock-oriented approach (that changes in stock prices cause changes in exchange rates) were tested using linear and nonlinear, asymmetric ARDL models. In the short-run, we found evidence favoring both models. In six of the G7 countries, exchange rate fluctuations significantly impact stock prices in the short-run and in five countries, stock prices significantly affect exchange rates. Thus, the foreign exchange and stock markets generally affect each other in the short-run. However, in the long-run, the flow-oriented model is not supported, while the stock-oriented, or portfolio balance approach receives considerable support. These results are generally consistent with the recent literature, except that we have provided stronger support for the stock-oriented (portfolio balance) approach, than in most other studies. By employing nonlinear, asymmetric ARDL models with a longer data series that includes many observations after the Global Financial Crisis, and allowing for structural breaks, we discover a long-run relationship from stock prices to exchange rates that is not spurious in four of the G7 countries (France, Germany, Italy, and the UK).

Specifically, changes in stock prices have both short-run and long-run effects on the exchange rates of France, Italy, and the UK, only short-run effects on the exchange rates of Canada, Japan, and the US, and only a long-run effect on the exchange rate of Germany. The results for Canada suggest that rising (falling) stock prices lead to short-run dollar appreciation (depreciation), which is consistent with the portfolio balance approach. Similarly, the finding for the UK that rising (falling) stock prices lead to long-run pound appreciation (depreciation) is consistent with the portfolio balance approach. For Italy, falling stock prices lead to currency depreciation in both the short-run and long-run, which is consistent with the portfolio balance approach. However, rising stock prices lead to currency depreciation in the long-run, which is inconsistent with the portfolio balance approach. Results for France and Germany that rising (falling) stock prices lead to currency depreciation (appreciation) in the long-run, is not consistent with the portfolio balance approach. In Japan, falling stock prices lead to short-run yen appreciation, which is again inconsistent with the portfolio balance approach. Also, rising (falling) US stock prices lead to dollar depreciation (appreciation) the dollar in the short-run, which is not consistent with the portfolio balance approach. Nevertheless, our general findings based on nonlinear ARDL models provide considerable support for the portfolio balance approach that stock prices affect exchange rates. Checking for robustness using VAR models and Granger causality tests indicates unidirectional causality from rising and/or falling stock prices to rising and/or falling exchange rates in all G7 countries, except Italy where the flow-oriented approach seems to be supported as the evidence suggest that currency appreciations and depreciations Granger cause falling stock prices.

Our findings should be important for both policy-makers and investors. The results indicate that the two markets affect each other in the short-run across all G7 countries. This implies that changes in either stock prices or exchange rates are transmitted quickly to the other market and that policymakers should be aware of this relationship. Furthermore, the adjustments to long-run equilibrium in France, Germany, Italy, and the UK seem to primarily occur through the stock markets. Given this outcome, and since that Granger causality tests indicate that stock prices Granger-cause exchange rates in all the countries except in Italy, our results suggest that policies that reduce stock market uncertainty may reduce exchange rate volatility in the G7 countries. Thus, stock market stabilization policies may help to stabilize foreign exchange markets. Furthermore, policymakers should recognize that domestic stock prices have a long-run impact on a country's exchange rate, but not vice-versa. Hence, attempting to manipulate exchange rates to boost stock markets in the long-run and/or using information available in foreign exchange market for long-run prediction of stock prices may not be an effective strategy. Finally, since changes in stock prices and exchange rates impact each other in the short-run, investors may be able to use the information available in one market to help predict the behavior of the other market for short-run hedging and speculation purposes.

## CRediT authorship contribution statement

**Salah A. Nusair:** Conceptualization, Methodology. **Dennis Olson:** Conceptualization, Writing – review & editing.

## Declaration of Competing Interest

The authors declare that they have no known competing financial interests or personal relationships that could have appeared to influence the work reported in this paper.

## Acknowledgment

We would like to thank an anonymous referee and the Editor of the journal for helpful and useful comments that improved the paper significantly. We are grateful to Mohsen Bahmani-Oskooee for his valuable input and useful discussion. The usual disclaimer applies.

## Appendix

### *Data Source*

All data are monthly. Stock price indices are closing prices and are extracted from Yahoo Finance. Nominal effective exchange rates are extracted from the Bank of International Settlements (https://www.bis.org/statistics/eer.htm). The following table provides the range of the sample for each country and the name of the stock market price index:

| Country | Sample | Stock market index name |
|---|---|---|
| Canada | 06 / 1979 to 02 / 2020 | S&P/TSX Composite |
| France | 03 / 1990 to 02 / 2020 | France CAC 40 |
| Germany | 12 / 1987 to 02 / 2020 | Frankfurt DAX |
| Italy | 12 / 1997 to 02 / 2018 | FTSE MIB |
| Japan | 06 / 1973 to 02 / 2020 | NIKKEI 225 |
| UK | 01 / 1984 to 02 / 2020 | FTSE 100 |
| USA | 06 / 1973 to 02 / 2020 | S&P 500 |

## References

Aggarwal, R., 1981. Exchange rates and stock prices: A study of the US. capital markets under floating exchange rates. Akron Bus. Econ. Rev. 12, 7–12.
Ajayi, R., Mougoue, M., 1996. On the dynamic relation between stock prices and exchange rates. J. Financ. Res. 19 (2), 193–207.
Bahmani-Oskooee, M., Mohammadian, A., 2018. Asymmetric effects of exchange rate changes on domestic output in emerging countries. Emerg. Mark. Finance Trade 54 (6), 1442–1459.
Bahmani-Oskooee, M., Saha, S., 2018. On the relation between exchange rates and stock prices: a nonlinear ARDL approach and asymmetry analysis. J. Econ. Finance 42, 112–137.
Bahmani-Oskooee, M., Saha, S., 2016. Do exchange rates have symmetric or asymmetric effects on stock prices? Global Finance J. 31, 57–72.
Bahmani-Oskooee, M., Saha, M., 2015. On the relation between stock prices and exchange rates: A review article. J. Econ. Stud. 42 (4), 707–732.
Bahmani-Oskooee, M., Sohrabian, A., 1992. Stock prices and the effective exchange rate of the dollar. Appl. Econ. 24 (4), 459–464.
Balke, N., Fomby, T., 1997. Threshold cointegration. Int. Econ. Rev. 38 (3), 627–645.
Banerjee, A., Dolado, J., Mestre, R., 1998. Error-correction mechanism tests for cointegration in a single-equation framework. J. Time Ser. Anal. 19 (3), 267–283.
Caporale, G., Hunter, J., Ali, F., 2014. On the linkages between stock prices and exchange rates: evidence from banking crisis of 2007–2010. Int. Rev. Financ. Anal. 33, 87–103.
Chen, S.-W., Chen, T.-C., 2012. Untangling the non-linear causal nexus between exchange rates and stock prices: new evidence from the OECD countries. J. Econ. Stud. 39 (2), 231–259.
Chkili, W., Neguyen, D., 2014. Exchange rate movements and stock market returns in a regime-switching environment: evidence for BRICS countries. Res. Int. Bus. Finance 31, 46–56.
Chow, E., Lee, W., Solt, M., 1997. The exchange rate risk exposure of asset returns. J. Bus. 70 (1), 105–123.
Dornbusch, R., Fisher, S., 1980. Exchange rates and the current account. Am. Econ. Rev. 70 (5), 960–971.
Engle, R., Granger, C., 1987. Co-Integration and error correction: Representation, Estimation, and Testing. Econometrica 55 (2), 251–276.
Frenkel, J., 1976. A monetary approach to the exchange rate: doctrinal aspects and empirical evidence. Scand. J. Econ. 78 (2), 200–224.
Granger, C., 1969. Investigating causal relations by econometric models and cross-spectral methods. Econometrica 37 (3), 424–438.
Gregory, A., Hansen, B., 1996. Residual-based tests for cointegration in models with regime shifts. J. Economet. 70, 99–126.
Johansen, S., 1998. Likelihood-based inference in cointegrated vector autoregressive models. Economet. Theory 14 (4), 517–524.
Kollias, C., Mylonidis, N., Paleologou, S.-M., 2012. The nexus between exchange rates and stock markets: evidence from the euro-dollar rate and composite European stock indices using rolling analysis. J. Econ. Finance 36, 136–147.
Kollias1, C., Papadamou, S., Siriopoulos, C., 2016. Stock markets and effective exchange rates in European countries: threshold cointegration findings. Eurasian Econ. Rev. 6, 215–274.
Koulakiotis, A., Kiohos, A., Babalos, V., 2015. Exploring the interaction between stock price index and exchange rates: an asymmetric threshold approach. Appl. Econ. 47 (13), 1273–1285.
Liang, C.-C., Lin, J.-B., Hsu, H.-C., 2013. Reexamining the relationships between stock prices and exchange rates in ASEAN-5 using panel Granger causality approach. Econ. Model. 32, 560–563.
Morley, B., Pentecost, E., 2000. Common trends and cycles in G-7 countries exchange rates and stock prices. Appl. Econ. Lett. 7 (1), 7–10.
Nieh, C., Lee, C., 2001. Dynamic relationship between stock prices and exchange rates for G-7 countries. Quart. Rev. Econ. Finance 41 (4), 477–490.
Pesaran, M., Shin, Y., Smith, R., 2001. Bound testing approaches to the analysis of level relationship. J. Appl. Economet. 16 (3), 289–326.
Ratner, M., 1993. A cointegration test of the impact of foreign exchange rates on US stock market prices. Global Finance J. 4 (2), 93–101.
Roll, R., 1992. Industrial structure and the comparative behavior of international stock market indices. J. Finance 47 (1), 3–41.
Salisu, A., Ndako, U., 2018. Modelling stock price-exchange rate nexus in OECD countries: a new perspective. Econ. Model. 74, 105–123.
Shin, Y., Yu, B., Greenwood-Nimmo, M.J., 2013. Modelling asymmetric cointegration and dynamic multipliers in a nonlinear ARDL framework. In: Horrace, W., Sickles, C., Robin, C. (Eds.), Festschrift in Honor of Peter Schmidt. Springer Science & Business Media, New York.
Tang, X., Yao, X., 2018. Do financial structures affect exchange rate and stock price interaction? Evidence from emerging markets. Emerg. Mark. Rev. 34, 64–76.
Wong, H., 2017. Real exchange rate returns and real stock price returns. Int. Rev. Econ. Finance 49, 340–352.
Xie, Z., Chen, S.-W., Wu, A.-C., 2020. The foreign exchange and stock market nexus: New international evidence. Int. Rev. Econ. Finance 67, 240–266.
Yang, S.-Y., Doong, S.-C., 2004. Price and volatility spillovers between stock prices and exchange rates: empirical evidence from the G-7 countries. Int. J. Bus. Econ. 3 (2), 139–153.