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Does foreign direct investment crowd in or crowd out private domestic investment in China? The effect of entry mode☆

George S. Chen[a,*], Yao Yao[b], Julien Malizard[c]

[a] UNE Business School University of New England, Room 809, W40, Armidale, NSW 2351 Australia
[b] UNE Business School University of New England, Room G03, W39, Armidale, NSW 2351 Australia
[c] UMR 5113 GREThA, France

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ABSTRACT

Using quarterly data spanning from 1994Q1 to 2014Q4, we find a neutral relationship between foreign direct investment (FDI) and domestic investment in China. However, when we consider the entry mode chosen by foreign investors, we find that whilst equity joint venture (EJV) crowds in domestic investment, wholly foreign-funded enterprise (WFFE) crowds it out. Our results remain robust under alternative estimators and across different time periods. Based on these results, we argue that the Chinese government needs to actively promote the formation of EJV and uses it as the catalyst for industrial upgrading in the economy.

1. Introduction

Following the famous 1992 Southern Tour by then the Chinese leader, Deng Xiaoping, foreign direct investment (FDI) in China enjoyed, for the most part of the last two decades, unprecedented growth. Driven by the renewed interest following its accession to the World Trade Organisation (WTO) in 2001, China surpassed the United States as the world's most popular destination among international investors for the very first time in 2003. To put this achievement into perspective, UNCTAD (2015) reported that in 2013 China hosted approximately US$1,085 billion or around 13% of the total FDI stock in the developing world. With this backdrop, many researchers argue that FDI plays a catalytic role in modernizing the Chinese economy, not only in terms of promoting technological transfers, but also in transforming business practices and the institutional environment in China (Berthélemy and Démurger, 2000; Cole et al., 2009; Elliott et al., 2013; Havrylchyk and Poncet, 2007; Hering and Poncet, 2010; Lo et al., 2016; Long et al., 2015; Madariaga and Poncet, 2007; Whalley and Xin, 2010; Yang et al., 2013). Undoubtedly, these positive spillovers further reinforce the willingness of the Chinese firms to re-invest, contributing to China's impressive growth record since the early 1990s. For example, Sun (1998) attributes at least one third of the growth in domestic private investment (henceforth, domestic investment) in 10 coastal provinces during the 1983–1995 period to the influx of FDI. Meanwhile, Xu and Wang (2007) examine China's national accounts data and find that FDI crowds in domestic investment from 1980 to 1999. Tang et al. (2008) extend this idea to show unidirectional causality running from FDI to domestic investment in China over the 1988–2003 period.

However, not everyone shares such an optimistic sentiment and argues that FDI crowds out domestic investment, damaging the long-term prosperity of the host country. In part, this pessimistic view stems from the premise that FDI intensifies competition in local factor and product markets that either reduces the willingness of the indigenous firms to re-invest or drives the incompetent ones out of business altogether (Aitken and Harrison, 1999). Meanwhile, Gall et al. (2014) show that a sudden withdrawal of FDI may severely hinder the growth prospect in those host countries with imperfect credit markets. Indeed, as witnessed during the 1997 Asian currency crisis, and more recently, the global financial crisis in 2008, an unexpected decline in FDI presented a major obstacle impeding the recovery process in many emerging economies. This crowding-out hypothesis echoes Huang's

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* Corresponding author.
E-mail addresses: george.chen@une.edu.au (G.S. Chen), yyao4@myune.edu.au (Y. Yao), julien.malizard@gmail.com (J. Malizard).

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(2003a, 2003b) argument that China's massive influx of FDI merely reflects institutional deficiency brought about by inefficient allocation of resources under a planned economy. Specifically, he suggests that preferential treatment given to state-owned enterprises by China's state-led banking sector has significantly limited the scope of growth in many private firms. In order to overcome this lack of credit availability and other finance constraints, many rapidly growing private firms start to seek foreign partners (Egger and Nelson, 2011). In this regard, FDI simply substitutes domestic investment, leaving little changes to the level of overall investment in China. This substitution hypothesis is supported by Braunstein and Epstein (2002), who examine the FDI–domestic investment nexus in 29 Chinese provinces during the 1986–1999 period and question the widely held belief that China's rapid ascendancy was largely propelled by FDI.

Despite a rather voluminous literature on the causes and effects of FDI in China, most studies have ignored the effect of entry mode on the relationship between FDI and domestic investment in the Chinese economy. According to China's National Bureau of Statistics (NBS), the top-three entry modes chosen by foreign investors are equity joint venture (EJV), contractual joint venture (CJV), and wholly foreign-funded enterprise (WFFE).[1] A recent study by Ashraf and Herzer (2014), who investigate the effect of entry mode on the FDI–domestic investment nexus in 100 developing countries, conclude that whilst there is a neutral relationship between merges and acquisitions (M & As) and domestic investment, WFFE tends to crowd out domestic investment. If their finding also holds true for China, then the policy makers need to curtail the growth of WFFE, which has been that most preferred entry mode in China since 1999, particularly among investors from Hong Kong, Macao, and Taiwan.

Our study represents the first systematic inquiry into the fundamental relationship between entry mode and domestic investment in China. Conceptually, we extend Ashraf and Herzer (2014) by also including EJV and CJV, two of the most popular entry mode chosen by foreign investors in China, in the analysis. Methodologically, our study examines the association between entry mode and domestic investment through the lens of the autoregressive distributed lag (ARDL) bounds test. According to Pesaran et al. (2001) and Narayan and Smyth (2005), this test delivers much better small sample properties and places less restrictive conditions on the order of integration in the model. Apart from the conceptual and methodological considerations, our analysis focuses on quarterly data spanning from 1994 to 2014, which takes into the account the impact of China's accession to the WTO in 2001 on the entry mode–domestic investment nexus.

In general, we find a neutral relationship between FDI and domestic investment in China for the entire sample period. However, when we introduce entry mode into the analysis, we find that EJV crowds in domestic investment, but WFFE crowds it out. Furthermore, we show that the nature of the FDI–domestic investment nexus changes over time. Specifically, we find that whilst FDI crowds in domestic investment prior to joining the WTO, FDI crowds it out during the post-WTO era. Based on these findings, we argue that the Chinese government should encourage the formation of EJV and uses it as the platform for encouraging industrial upgrading in the economy.

The remaining part of this paper is structured as follows. Section 2 provides a brief literature review on the current state of research on the FDI–domestic investment nexus. Section 3 describes the emerging trend of entry mode in China and argues for its inclusion in the analysis. The econometric framework and results are discussed in Sections 4 and 5, respectively. Section 6 concludes.

2. Literature review

2.1. Panel studies

In a recent seminal paper, Agosin and Machado (2005) apply the difference generalized method of moment (GMM) estimator to examine the effects of FDI on domestic investment in 36 developing countries during the period 1971–2000. In order to mitigate aggregation bias, they split these countries equally into 12 countries in each of the Asian, African and Latin American regions and find that FDI either exerts no influences over, or partially crowds out, domestic investment in the host country. Based on this finding, they challenge the notion that positive externalities brought about by FDI stimulate domestic investment in the host country and conclude that *"the effects of FDI on domestic investment are by no means always favourable, that simplistic policies towards FDI are unlikely to be optimal and, foremost, that more attention needs to be paid to economic policies that foster the domestic component of total investment"* (Agosin and Machado, 2005, p. 149).

In a following-up study, Morrissey and Udomkerdmongkol (2012) improve on Agosin and Machado (2005) by including governance as one of the control variables and apply the system GMM estimator to a panel consisting of 46 developing countries from 1996 to 2009.[2] In general, they find that FDI not only crowds out domestic investment in the host country, but the extent of such crowding out increases with better governance. In part, they attribute this finding to the fact that whilst good governance promotes FDI, it also creates fierce competition in the factor and product markets that reduces the willingness of inefficient indigenous firms to re-invest. Since domestic investment is often regarded as an engine of sustainable economic development, they share the view expressed by Alguacil et al. (2011) and suggest that *"policies designed to attract FDI are not sufficient to ensure economic growth"* (Morrissey and Udomkerdmongkol, 2012, p. 443).

Despite the attempt by Morrissey and Udomkerdmongkol (2012) to address various shortcomings in Agosin and Machado (2005), Farla et al. (2016) question the validity of the unfavorable findings against FDI in the host country. Conceptually, they criticize Morrissey and Udomkerdmongkol (2012) for using inappropriate proxies of foreign and domestic investment in the analysis that introduces downward bias on the estimates. Methodologically, this downward bias is further exacerbated by the fact that Morrissey and Udomkerdmongkol (2012) overlook the problem of instrument proliferation in their system GMM estimations. Applying properly specified system GMM estimator to the original Morrissey-Udomkerdmongkol dataset, they find that FDI crowds in domestic investment in the host country and conclude that *"policy aimed at stimulating FDI inflow is likely to have a positive effect on developing countries' economy"* (Farla et al. 2016, p. 7).

An important point demonstrated by Farla et al. (2016) is that the nature of the FDI–domestic investment nexus can be extremely sensitive to model specifications and prone to aggregation bias. In the case of the Morrissey-Udomkerdmongkol dataset, a potential source of aggregation bias can be traced to the mixed collection of developing countries at various stages of economic development. In theory, this mixed collection violates the homogeneity assumption imposed on the coefficients of the lagged dependent variables by the GMM estimators, *"when in fact the dynamics are heterogeneous across the panel"* (Herzer et al., 2008, p. 796). In order to mitigate aggregation bias, Ndikumana and Verick (2008) focus on 38 sub-Saharan African countries for the period 1970–2005 and find a

[1] The remaining entry modes reported by the NBS include joint exploration and FDI shareholding. In general, joint exploration is more common among natural resource-seeking foreign investors. Meanwhile, FDI shareholding usually involves a much larger minimum registered capital threshold and requires the Chinese entity to be divided into local and foreign shareholding, each with an equal par value (Wei et al., 2005). Our study excludes these two entry modes as they account for less than 5% of total registered FDI stock in China.

[2] Since Arellano and Bover (1995) show in their Monte Carlo simulations that lagged levels are often poor instrument for first differences, Agosin and Machado's (2005) choice of the difference GMM estimator may not be an appropriate choice.

crowding-in effect in this region. This result is further confirmed by Adams (2009), who examine 42 sub-Saharan African countries over a shorter sample period from 1990 to 2003. Meanwhile, some researchers manage aggregation bias by categorizing countries according to their stage of economic development. For example, Wang (2010) examines 50 countries over the period of 1970 to 2004 and finds that the cumulative effect of FDI on domestic investment is neutral in developed countries, but positive in less developed countries. Similarly, Al-Sadig (2013) divides a sample of 91 developing countries during the 1970–2000 period and find support for the crowding-in hypothesis, albeit the extent of this effect depends crucially on the availability of human capital in low-income countries.

2.2. Panel cointegration and causality studies

Recent advancements in econometric techniques have allowed researchers to take into account heterogeneity in cointegrated panels when uncovering the long-run relationship between FDI and domestic investment. For example, Jain et al. (2014) apply the panel fully-modified OLS (PFMOLS) and simple vector autoregressions (VAR) to test the causal effects between FDI and domestic investment in 22 emerging economies. By taking into account heterogeneous cointegrated panels, Jain et al. (2014, p. 1) find a positive and bidirectional FDI–domestic investment nexus in Asia for the period 1995–2007 and suggest that this result "*is consistent with the complementary hypothesis of neoclassical macroeconomic growth model in which it is often thought that FDI inflows complement the domestic investment.*" Based on a similar growth model framework, Omri and Kahouli (2014) examine 13 countries in the Middle East and North Africa during the period 1990–2010. Under a simultaneous-equations setup, they find a unidirectional causal relationship running from FDI to domestic capital for the region as whole. Similar to Alfaro et al. (2006), they suggest that improving domestic conditions not only attracts FDI, but also enables the host country to fully internalize positive spillovers brought about by FDI.

However, Strauss and Wohar (2004) suggest that the results from panel cointegration and causality studies may not be reliable if the panel is mixed with cointegrated and non-cointegrated relationships. Indeed, Banerjee et al. (2004) warn that inferences in these studies need to be interpreted with caution and suggest that analyzing time series for individual countries may be more appropriate.

2.3. Time series studies for individual countries

In examining the causal relationship between FDI and domestic investment in an individual country, two broad methodologies have emerged in this strand of literature with the aim of controlling endogeneity in FDI; namely, the error-correction model (ECM) and the ARDL model. For example, Kim and Seo (2003) apply the innovation accounting techniques in an unrestricted VAR system to the Korean data during the period 1985–1999 and find that while growth in domestic investment causes a fall in FDI, there exists a strong crowding-in effect prior to the onset of the Asian financial crisis. They attribute these results to dynamic endogeneity of FDI, which has been largely neglected in previous studies. In China, Tang et al. (2008) use a VAR system with ECM to find a crowding-in effect, but only unidirectional causality running from FDI to domestic investment. In Malaysia, Ang (2009b) divides domestic investment into private domestic investment and public investment, before fitting them to a vector ECM (VECM). During the period 1960 to 2003, he finds support for the crowding-in hypothesis. Similarly, Ghazali (2010) examine the Pakistani data spanning from 1981 to 2008 to find a crowding-in effect between FDI and net domestic real investment and bidirectional causality between these two types of investment. Meanwhile, using the firm-level data from Germany during the period 1991–2003, Arndt et al. (2010) find crowding in between inward FDI and domestic capital stock, but crowding out between outward FDI and domestic capital stock under a VECM representation. In a multi-country setting, when Qi (2007) applies the ECM to 47 countries during the period 1970–2003, he finds unidirectional causality running from domestic investment to FDI in developed countries, but bidirectional causality in developing countries. Furthermore, whilst the sign of the causal effect between FDI and domestic investment remains positive for developed countries, it is sometimes negative in developing countries. In part, he attributes these contradicting results to the country's oil-exporting status.

Based on the ARDL model, Goh and Wong (2014) consider the effect of inward and outward FDI on domestic investment in Malaysia from 1991Q1 to 2010Q3 and find support for crowding-in between inward FDI and domestic investment, but crowding-out between outward FDI and domestic investment. Furthermore, they argue that since the inward FDI–domestic investment nexus is relatively elastic compared to its outward FDI–domestic investment counterpart the Malaysian government should attract inward FDI and use it to offset the crowding-out effect brought about by outward FDI. Focusing on the FDI–domestic investment nexus in agriculture, Djokoto et al. (2014) also finds a crowding-in effect when the ARDL model is applied to Ghana from 1976 to 2007.

Thus far, our discussion suggests that the nature of the FDI–domestic investment nexus in the host country is extremely sensitive to the chosen methodology, time period, and aggregation bias. Nevertheless, the extant literature largely supports the crowding-in hypothesis, particularly in developing countries. However, after studying a similar nexus in Canada from 1948 to 1966, Van Loo (1977, p. 481) suggests that *"policy makers should not simply assume that foreign direct investment increases the ability of the economy to produce. Rather, it is important to examine precisely how those foreign funds are used."* To date, Ashraf and Herzer (2014) provides the only study that explicitly takes into account the entry mode of FDI in the host country. Specifically, they investigate the effect of WFFEs and M & As on domestic investment in 100 developing countries from 2003 to 2011 and find that whilst there is a neutral relationship between M & As and domestic investment, WFFEs tends to crowd out domestic investment. This result challenges the conventional wisdom that FDI generates a positive effect in the host country. In the spirit of Ashraf and Herzer (2014), this study sets out to investigate the effect, if any, entry-mode choice impacts on domestic investment in China.

3. Entry mode in China

As discussed in Section 1 the rise to prominence of FDI in China is not an entirely new phenomenon. However, what is new is that the preferred entry mode has significantly changed over the years. According to the National People's Congress (1979, 1986, 1988) and the Ministry of Foreign Trade and Economic Cooperation (1995), each entry mode is subject to different rules and regulations. Formally, an EJV is defined as a limited liability, legal entry where the management of routine operation, distribution of profit and commitment of resources are based on the equity contribution by the Chinese and foreign partners. In contrast, a CJV represents a more flexible entry mode, where parties to the transaction are free to negotiate on the organizational structure, with the rights and obligations clearly set out in the contract (Wei et al., 2005). Finally, a WFFE is a limited liability legal entity solely owned and operated by foreign investors.

The choice of entry mode can reflect foreign investor's motives for operating beyond national boundaries. Given the varying degree of control, resource commitment, and risk exposure, each entry mode interacts differently with its host economy (Blomström et al., 2001; Canabal and White Iii, 2008; Meyer et al., 2009). For instance, EJV and CJV create an ideal environment for their local Chinese partners to learn the best practice from their foreign counterparts (Filatotchev et al., 2007; Tse et al., 1997). This positive spillover effect is not restricted to the transfer of technological expertise, but also in managerial and marketing known-how. Indeed, this point corresponds to Pomfret's (1991, p. 135) assertion that *"what was missing in PRC, rather than capital, was the knowledge of how to make bags or teddy bears or wind-up pandas or cigarette lighters in attractive designs or reasonable quality standards and of how to market them overseas."*.[3]

With these definitions of entry mode, Fig. 1 shows that while EJV, CJV, and WFFE collectively accounted for over 97% of annual FDI inflow during the 1995–2014 period the share of CJV shrunk from 20% in the beginning to less than 1% in the end. In relation to this, in 1999 WFFE replaced EJV as the most preferred entry mode. In part, these changes reflected significant improvements in China's legal framework and institutional environment since the turn of this century (Chen, 2011). And partly, more than 20 years of experience in navigating the Chinese business scene encouraged many foreign investors to establish and operate WFFEs independently (Wei et al., 2005).

From the outset, both EJV and CJV seem to crowd in domestic investment as they require their Chinese partners to also contribute capital in the newly established joint venture. In contrast, the lack of the Chinese involvement in WFFE may suggest that FDI crowds out domestic investment in China. However, such a conjecture may be premature as it overlooks the fact that many WFFEs in China are large in scale and engage in capital-intensive industries (Buckley et al., 2007; Huang, 2004; Zhang, 2005). These characteristics suggest that WFFEs could exert significant spillovers on industry linkages and competition in the host country (Agarwal and Ramaswami, 1992). To visualize this, suppose that a WFFE has developed close ties with numerous indigenous firms, either as an upstream supplier or a downstream customer. These ties, in turn, create substantial transfers of tacit and non-tacit knowledge within the local supply chain, thereby, creating a fertile environment for those indigenous firms to develop and expand (Blomström et al., 2001). In this scenario, this WFFE crowds in domestic investment. However, if the primary objective of the WFFE is to penetrate China's mass virgin market, then fierce competition is likely to drive out those inefficient Chinese firms over time. In this regard, WFFE crowds out domestic investment.

In short, our preceding discussion suggests that entry mode can be an important determinant on the nature of the FDI–domestic investment nexus in China. Any study that fails to explicitly consider the role of entry mode is prone to aggregation bias and runs the risk of arriving at erroneous policy implications.

4. Econometric framework

4.1. Model specifications

We begin by not distinguishing the entry modes chosen by foreign investors in China. Formally, the baseline model is given by Eq. (1) below:

$$DI_t = \kappa_1 + \alpha_1 FDI_t + X'\beta + \varepsilon_t \tag{1}$$

where DI_t and FDI_t, respectively, refer to the level of domestic investment and FDI in quarter t, X' represents a vector of control variables, κ_1 is the constant term, and ε is the stochastic error term. In terms of the control-variable set, we use output (*GDP*) to capture the positive relationship between economic growth and fixed capital formation induced by strong expected sales, cash flows, and profitability (Ashraf and Herzer, 2014; Mody and Murshid, 2005; Tang et al., 2008). Meanwhile, we measure the opportunity cost of investment by the one-year benchmark lending rate (*RATE*) (Ashraf and Herzer, 2014; Wang, 2010). Given the significance of the exporting sector to the Chinese economy, we also include the volume of export (*EX*) in the model (Adams, 2009; Ashraf and Herzer, 2014; Wang and Wang, 2015; Wang and Wong, 2009; Yang and Mallick, 2014). Except for the lending rate, we select the log level form of the variables in Eq. (1) during estimations as it exhibits greater variations than their respective ratio form (Tang et al., 2008). For this study, we are interested in the sign and magnitude of α_1 in Eq. (1). Specifically, a positive and statistically significant α_1 indicates a crowding-in effect between FDI and domestic investment, a negative and statistically significant α_1 lends support to the crowding-out hypothesis (see, for example, Agosin and Machado, 2005; Farla et al., 2016; Morrissey and Udomkerdmongkol, 2012).

However, as discussed in Section 3, the relationship between FDI and domestic investment may vary depending on the entry mode in question. This distinction is important because it is possible for a crowding-in effect between EJV or CJV and domestic investment to be offset by crowding out between WFFE and domestic investment, leading to an overall, neutral FDI–domestic investment nexus. Put differently, since α_1 in Eq. (1) measures the average impact of over the three entry modes on domestic investment, it cannot inform policy makers on the ideal entry model to promote. To avoid this aggregation bias, we replace FDI in Eq. (1) with EVJ, CJV and WFFE. This augmented model is given in Eq. (2) below:

$$DI_t = \kappa_2 + \delta_1 EJV_t + \delta_2 CJV_t + \delta_3 WFFE_t + X'\beta + \varepsilon_t \tag{2}$$

Except for EJV_t, CJV_t, and $WFFE_t$, which refer to the level of FDI associated with each entry mode in year t, all other variables in Eq. (2) follow the same definition as in Eq. (1). Once again, a positive and statistically significant δ_i indicates crowding in between entry mode i and domestic investment in China. In contrast, a negative and statistically significant δ_i represents a crowding-out effect.

4.2. The ARDL bounds test

In time series analysis, a meaningful cointegrated relationship must exist in the model in order to rule our spurious results. In the extant literature, both bivariate and multivariate cointegration tests are often used to identify the long-run relationship among key variables of interest (Engle and Granger, 1987; Johansen, 1988, 1991; Johansen and Juselius, 1990). In theory, the multivariate cointegration test seems to be more efficient as it uncovers multiple cointegrating vectors. However, Ang (2009a) suggests that such test could be hard to interpret if more than one cointegrating vector is found in the model. In the extreme case of a model with mixed orders of integration, neither bivariate nor multivariate cointegration test is appropriate.

An alternative to these aforementioned cointegration tests is the ARDL bounds test postulated by Pesaran and Pesaran (1997), Pesaran and Smith (1998), and Pesaran et al. (2001), among others. This approach has four main advantages over conventional cointegration tests. First, it can be applied to models with mixed order of integration. Second, it exhibits superior small-sample property than conventional cointegration tests (Narayan, 2005; Narayan and Smyth, 2005; Smyth and Narayan, 2015). Third, a correctly specified lag structure not only controls for serial correlation, but also minimizes potential endogeneity in the model (Pesaran and Shin, 1998). Fourth, and perhaps more importantly, we can obtain a dynamic unrestricted ECM (UECM) by applying simply linear transformation to the specified ARDL model, with the transformed UECM enjoying the benefit of combining short-run dynamics and long-run equilibrium together without losing any significant information (Baek, 2016; Sbia et al., 2014). In our study, the corresponding UECM of Eqs. (1) and (2) can be expressed as follows:

$$\begin{aligned}\Delta DI_t = {} & \chi_1 DI_{t-1} + \chi_2 FDI_{t-1} + \chi_3 GDP_{t-1} + \chi_4 EX_{t-1} + \chi_5 RATE_{t-1} \\ & + \sum_{i=1}^{p} \phi_i \Delta DI_{t-i} + \sum_{j=0}^{q} \phi_j \Delta FDI_{t-j} + \sum_{k=0}^{r} \phi_k \Delta GDP_{t-k} + \sum_{l=0}^{s} \phi_l \Delta EX_{t-l} \\ & + \sum_{m=0}^{n} \phi_l \Delta RATE_{t-m} + \varepsilon_t\end{aligned} \tag{3}$$

[3] PRC stands for the People's Republic of China.

$$\Delta DI_t = \eta_1 DI_{t-1} + \eta_2 EJV_{t-1} + \eta_3 CJV_{t-1} + \eta_4 WFFE_{t-1} + \eta_5 GDP_{t-1} + \eta_6 EX_{t-1} + \eta_7 RATE_{t-1} + \sum_{i=1}^{p} \lambda_i \Delta DI_{t-i} + \sum_{j=0}^{q} \lambda_j \Delta EJV_{t-j} + \sum_{n=0}^{u} \lambda_n \Delta CJV_{t-n} + \sum_{o=0}^{v} \lambda_o \Delta WFFE_{t-o} + \sum_{k=0}^{r} \lambda_k \Delta GDP_{t-k} + \sum_{l=0}^{s} \lambda_k \Delta EX_{t-l} + \sum_{m=0}^{n} \lambda_k \Delta RATE_{t-m} + \varepsilon_t \tag{4}$$

where Δ denotes the first-difference operator. Since our sample size is relatively small, we employ the Akaike information criterion (AIC) to select the optimal lag structure in Eqs. (3) and (4) (Lütkepohl, 2005). In addition, we perform the usual diagnostic tests for serial correlation, functional form, normality, and heteroskedasticity to ensure correct model specification for both equations. Finally, we test the stability of our estimates based on the cumulative sum (CUMSUM) and cumulative sum of squares (CUSUMSQ) statistics, which are updated recursively and plotted against the break points (Brown et al., 1975; Pesaran et al., 2001).

Essentially, the ARDL bounds test involves testing the null hypothesis of no cointegrated relationship in Eq. (3) (H_0: $\chi_1 = \cdots = \chi_5 = 0$)against the alternative hypothesis of a cointegrated relationship(H_1: $\chi_1 \neq \cdots \neq \chi_5 \neq 0$). Similarly, we test the null and alternative hypotheses for Eq. (4) as H_0: $\eta_1 = \cdots = \eta_7 = 0$ versus H_1: $\eta_1 \neq \cdots \neq \eta_7 \neq 0$. Since these hypothesis tests are effectively testing the joint significance of coefficients, we compute the *F*-statistic and compare it against the upper and lower bound critical value provided by Pesaran et al. (2001). It is worth pointing out that the computed *F*-statistic follows a non-standard distribution, which depends on sample size, degree of freedom, order of integration, and the inclusion of an intercept and/or a trend (Narayan and Narayan, 2005). As expected, we reject H_O if the *F*-statistic exceeds the upper-bound critical value. Meanwhile, we cannot reject H_O if the *F*-statistic falls below the lower-bound critical value. However, the test becomes inconclusive if the *F*-statistic lies within the range between the lower and upper bounds.

Once our variables are found to be cointegrated, we can obtain the ARDL long-run estimates in Eqs. (3) and (4). In order to ensure the robustness of our results, we consider three alternative estimators; namely, full-modified OLS (FMOLS), canonical correlation regression (CCR), and dynamic OLS (DOLS). We choose these estimators on the basis that: (*a*) FMOLS provides reliable point estimates and test statistics in a finite sample; (*b*) CCR eliminates asymptotically the endogeneity in the model; and (*c*) DOLS produces a richer pattern of data generation process as it takes into account both the leads and lags of the variables in the model (Narayan and Narayan, 2005; Park, 1992; Phillips and Loretan, 1991).

4.3. Data

We use quarterly data spanning the period 1994Q1–2014Q4 to study the FDI–domestic investment nexus in China. Specifically, we obtain statistics on FDI inflow and entry mode from China's Ministry of Commerce and figures on *GDP* and gross fixed capital formation (GFCF) from the CEIC database. According to Agosin and Machado (2005), given that FDI is recorded under the financial balance of payments and GFCF as a part of a country's national account, these two series are reported under different frameworks. This distinction is important because GFCF effectively measures the increase in the existing productive capacity, but FDI may or may not expand productive capacity. For instance, M & As often result in the transfer of ownership that need not necessarily affect existing productive capacity. In contrast, greenfield FDI often enlarges the existing pool of fixed capital stock that raises productive capacity in the future. In this study, we construct the DI series by subtracting FDI from GFCF, on the basis that Fig. 1 and OECD (2006) show that greenfield FDI (in the form of WFFE) accounted for the majority of FDI in China in recent years.

We then collect the data on export volume and the nominal lending from the IMF and the Federal Reserve of St. Louis, respectively.[4] Wherever appropriate, we convert the US dollar-denominated series into the Chinese currency using the quarterly average exchange rate published by the IMF and convert the nominal variables into their real counterparts by the GDP deflator released by China's National Statistical Bureau. With the exception of the real lending rate, we transform all variables into natural logarithm to eliminate complications arising from different units of measurement in the variables.

Finally, a visual inspection of the time series plot reveals strong seasonality patterns in our variables (Rawski, 2002; Tang et al., 2008). Following Johansson (2009), we deseasonalize our variables by the US Census Bureau's X-12 filter and present their descriptive statistics in Table 1.

5. Results

5.1. The Clemente-Montanes-Reyes unit root test

While the ARDL bounds test places fewer restrictions on the order of integration than conventional cointegration tests, it can still be biased if any of the variables in the model is integrated of order two or higher. As such, we carry out the ADF and DF-GLS unit root tests on the variables and conclude that none of them is integrated of order higher than one.[5] However, since both tests fail to take into account possible structural breaks, we also perform the Clemente-Montanes-Reyes (CMR) (1998) detrended structural break unit root test. This approach follows an additive outlier model by plugging out sudden changes in the mean of the variable as well as gradual changes in the mean of the variables tested by innovative outlier. Table 2 presents the CMR results for one or two structural breaks in the series. With the exception of the lending rate (*RATE*) we conclude that all variables are I(1). Notably, structural breaks for many variables at level occurred around 2003 and 2008, corresponding to the China's accession to the WTO and the fallout from GFC.[6]

5.2. The ARDL bounds test

Having confirmed that all of our variables are integrated of either I(0) or I(1), we now perform the ARDL bounds test. Table 3 reports the test results for Eqs. (3) and (4) and the relevant diagnostic tests with respect to their own UECM. Specifically, when comparing the *F*-statistic in Panel I of Table 3 to the critical values provided by Pesaran (2001), we reject the null hypothesis and conclude that there is a cointegrating relationship between FDI and domestic investment in China over the period 1994–2014. We reach the same conclusion when replacing FDI with different entry modes. Our results are validated by a battery of diagnostic tests reported in Panel II of Table 3. For example, the Jarque-Bera (J-B), ARCH LM and Breusch-Godfrey (B-G) LM tests all conclude that the stochastic error term in Eqs. (3) and (4) is normally distributed, homoscedastic, and not serially correlated. Meanwhile, the CUSUM and CUSUMSQ tests suggest that the underlying cointegrated relationship is stable over the sample period.[7]

Since we discovered structural breaks in some of the series in Section 5.1, it may undermine the reliability of our bounds test results

[4] We estimate the real lending rate by subtracting inflation from the nominal lending rate, on the basis that inflation and the nominal interest rate in China remained relatively stable throughout the sample period.

[5] The ADF and DF-GLS results are not reported here to conserve space but are available upon request from the authors.

[6] Although China was officially granted a WTO membership in December, 2001, it was allowed to take an incremental approach in fulfilling its membership obligations. As a result, the effect of the WTO accession did not occur after three to five years from 2001.

[7] To conserve space, we do not report the CUSUM and CUSUMSQ tests. These results are available upon request from the authors.

Table 1
Descriptive statistics.

	DI	*FDI*	*CJV*	*EJV*	*WFFE*	*GDP*	*EX*	*RATE*
Mean	15.508	12.453	9.500	11.152	11.876	15.294	13.907	4.395
Median	15.470	12.475	9.441	11.178	12.094	15.194	14.087	3.250
Maximum	17.802	13.028	10.771	11.409	12.7811	16.578	15.140	10.440
Minimum	12.209	11.865	8.287	10.813	10.379	13.834	12.474	2.700
Std. dev	1.430	0.379	0.739	0.142	0.726	0.788	0.903	2.485
Skewness	0.015	0.019	0.152	−0.166	−0.444	0.070	−0.142	1.589
Kurtosis	1.823	1.524	1.439	2.116	1.819	1.804	1.425	3.833
J-B test	4.740	7.450	8.641	3.0471	7.467	4.950	8.751	36.857
p-value	(0.094)	(0.024)	(0.013)	(0.218)	(0.024)	(0.084)	(0.013)	(0.000)

Note: all statistics are calculated from seasonally-adjusted series except the *RATE*.

Table 2
The Clemente-Montanes-Reyes structural break unit root test.

	Innovative outliers						Additive outliers					
	Level			First-differenced			Level			First-differenced		
	T_{B1}	T_{B2}	Test statistic	T_{B1}	T_{B2}	Test statistic	T_{B1}	T_{B2}	Test statistic	T_{B1}	T_{B2}	Test statistic
FDI	2001q4		-3.046(0)	2007q4		-7.192(3)***	2003q2		-3.000(0)	2007q3		-8.597(1)***
	2001q4	2006q2	-4.712(0)	2002q4	2007q4	-11.854(0)***	2001q3	2007q2	-4.214(2)	2002q3	2007q3	-9.525(1)***
EJV	2009q1		-2.346(4)	2007q4		-7.051(3)***	2001q3		-2.888(1)	2007q3		-6.304(3)***
	1997q4	2009q3	-4.844(0)	2002q4	2007q4	-11.921(0)***	1998q1	2010q2	-4.626(0)	2002q2	2007q3	-8.295(1)***
CJV	2003q1		-2.719(4)	2011q4		-7.567(3)***	2004q2		-1.835(4)	2011q3		-7.024(3)***
	1993q3	2004q3	-3.936(4)	2011q4	2012q4	-7.020(3)***	2000q2	2004q2	-3.105(4)	1998q2	2011q3	-7.642(3)***
WFFE	2000q2		-2.794(0)	2007q4		-7.111(3)***	2003q2		-2.693(0)	2007q3		-5.342(2)**
	2000q2	2006q2	-3.322(0)	2002q4	2007q4	-8.013(3)***	2001q3	2007q2	-3.842(0)	2002q3	2007q3	5.706(2)**
GDP	2004q3		-2.036(4)	1996q1		-8.484(0)***	2006q4		-2.447(0)	1996q1		-8.512(0)***
	2003q2	2006q2	-2.763(4)	1996q3	2003q1	-9.524(0)***	2005q1	2010q3	-3.064(1)	1996q4	2002q2	-9.890(0)***
EX	2001q3		-3.184(4)	2008q4		-5.630(3)***	2005q1		-2.463(0)	2008q3		-4.456(3)**
	2001q3	2008q4	-3.493(4)	2008q2	2008q4	-6.636(3)***	2004q1	2008q4	-2.943(0)	1999q3	2008q3	-6.636(1)**
RATE	1997q3		-7.963(0)***	1999q1		-9.491(0)***	1998q2		-4.722(2)**	1997q3		-6.266(1)***
	1995q4	1997q3	-9.652(1)***	1997q3	1999q2	-8.247(3)***	1996q3	1998q2	-2.800(5)	1997q2	1998q4	-6.639(2)**
DI	2002q2		-2.907(5)	2003q4		-10.676(1)***	2005q1		-2.271(5)	1995q2		-6.404(2)***
	2002q2	2008q2	-3.747(5)	2003q4	2008q4	-11.536(1)***	2004q2	2010q1	-4.401(0)	2003q3	2008q3	-9.552(1)***

Note: The asterisk ***, **, * denote the significance at 1%, 5% and 10% levels, respectively. T_{B1} and T_{B2} refer to the dates of the structural break. Optimal lag lengths are reported in parentheses. The critical value for the one and two structural break tests are taken from Perron and Vogelsang (1992) and Clemente et al. (1998), respectively.

obtained thus far (Sbia et al., 2014). To ensure the robustness of our results, we also report the Gregory and Hansen (1996a; 1996b) structural break cointegration test in Table 4.[8] In general, the Gregory–Hansen tests reject the null hypothesis of no cointegrated relationship in Eqs. (3) and (4).

5.3. The ARDL estimates

Given that cointegration exists in Eqs. (3) and (4), we now move to estimate the long-run FDI–domestic investment and entry mode–domestic investment nexuses in China. Since the coefficient of FDI for Eq. (3) in Panel III of Table 3 is statistically insignificant, we conclude that there is a neutral FDI–domestic investment nexus in China. This is in direct contrast to Xu and Wang (2007) and Tang et al. (2008), who report a crowding-in nexus between these two types of investment. We offer several plausible explanations for these mixed findings. First, our study covers both the pre- and post-WTO periods, whereas Xu and Wang (2007) and Tang et al. (2008) only focus on the former. As Chen (2011) point out the FDI regime was greatly liberalized by the Chinese government during the post-WTO period. However, this more liberal institutional environment also lured many market-seeking foreign investors with advanced technologies and marketing networks that drove many inefficient Chinese firms out of the market. Second, the neutral FDI–domestic investment nexus can be explained by the change in the preferred entry mode from EJV to WFFE in recent years, reflecting the countervailing effects of EJV and WFFE on domestic investment discussed in Section 3. Last, but not the least, these mixed findings lend support to Farla et al. (2016), who highlight the sensitive nature of the relationship between FDI and domestic investment to the chosen estimation methodology. Indeed, we will examine the robustness of our results in Section 5.4.[9]

Indeed, the ARDL estimates of Eq. (4) support the view that entry mode affects the FDI-domestic investment nexus in China. Panel III of Table 3 shows that a 1% increase in WFFE reduces domestic investment by 0.61%, all things being equal. Huang (2003b) and Wei et al. (2005) attribute this crowding-out relationship to more efficient WFFEs displacing their inefficient indigenous counterparts. This, coupled with preferential treatment and fiscal concessions to WFFEs from the Chinese government, deters many indigenous firms from expanding their existing production capacity. Over time, this reluctance causes FDI to crowd out domestic investment in China.

Surprisingly, we find that EJV crowds in, but CJV crowds out, domestic investment in China. To be more precise, Panel III of Table 3 shows that a 1% increase in EJV raises domestic investment by 0.55%, but a 1% increase in CJV reduces domestic investment by 0.35%. In

[8] Table 4 examines four different settings in the Gregory–Hansen structural break cointegration test, including the break in constant, break in constant and trend, break in constant and slope, and break in constant, slope, and trend.

[9] Following the recommendation from the anonymous referee, we also apply the Johansen cointegration test presented by Tang et al. (2008). We find that the Johansen cointegration test results are qualitatively the same as our ARDL bounds test during our sample period. Results are available upon request.

Table 3
The ARDL bounds test and long-run estimates.

	Baseline	Augmented
	Eq. (3)	Eq. (4)
Functional form	$F(DI \mid FDI;GDP;EX;RATE)$	$F(DI \mid EJV;CJV;WFFE;GDP;EX;RATE)$
Panel I: The bounds test		
Optimal lag structure[a]	4; 6; 3; 0; 6	8; 8; 4; 7; 7; 3; 6
F-statistic[b]	10.686***	11.403***
Panel II: Diagnostic tests		
R^2	0.682	0.942
Adj-R^2	0.541	0.825
F-statistic (UECM)	4.849***	7.999***
J-B Normality test	2.233	4.724
B-G LM test	[1] 1.263; [2] 1.277	[1] 0.219; [2] 0.945
ARCH LM test	[1] 0.158; [2] 0.432	[1] 0.052; [2] 1.510
Ramsay test	[1] 6.101**	[1] 4.460**
CUSUM	Stable	Stable
CUSUMSQ	Stable	Stable
Panel III: ARDL long-run estimates		
FDI	-3.963(3.160)	
EJV		0.545(0.125)***
CJV		-0.352(0.066)***
WFFE		-0.610(0.145)***
GDP	2.794(1.056)***	1.715(0.078)***
EXPORT	1.110(0.725)	0.255(0.077)***
RATE	0.217(0.165)	-0.036(0.014)**
Constant	4.867(13.175)	-9.982(1.129)***

Note: The asterisks ***, ** and * denote the significance at 1%, 5% and 10% levels, respectively. [a] The optimal lag structure of the ARDL models are determined by AIC. [b] The F-statistic is compared to the critical bounds computed by Pesaran et al. (2001) for restricted intercept and no trend. The brackets [ ] denotes the order of the diagnostic test. Standard error of the ARDL estimates are reported in the parentheses.

Table 4
The Gregory-Hansen structural break cointegration test.

Settings	Break in constant term	Break in the constant and trend	Break in constant and slope	Break in constant, slope and trend
Eq. (1)				
ADF test	-12.86(0)***	-13.60(0)***	-12.97(0)***	-12.78(0)***
TB	2002q4	2002q4	2007q3	2007q3
Eq. (2)				
ADF test	-14.19(0)***	-15.40(0)***	-12.37(0)***	-14.32(0)***
TB	2002q2	2002q4	2007q1	2007q4

Note: The asterisks ***, ** and * denote the significance at 1% level. The optimal lag length, which is reported in the parentheses, is selected by *t*-statistic method. The trimming region is set at 0.01. TB refers to the structural break date.

part, this conflicting result may be due to the inherent differences in these two forms of joint ventures. For example, the long investment horizon of an EJV encourages the Chinese and foreign partners to cooperate closely and expand their scale of operation, adding to domestic investment. Furthermore, some ex-EJV workers might start their own businesses that contributes to the pool of domestic investment. In contrast, the investment horizon of a typical CJV is usually shorter with limited participation by foreign investors. In fact, it is not uncommon for foreign investors in a CJV to pull out within a few years on an accelerated dividend payment schedule, before handing over all assets and control to their Chinese partners. Consequently, Huang (2003b) argues that CJVs are often set up by cash-strapped privately-owned Chinese firms as an alternative to accessing credit from the local banking sector. If this argument were correct the injection of FDI merely fills the shortfall in operating capital rather than being used for expanding production capacity, leaving domestic investment unchanged. Meanwhile, Fu (2008) suggests that CJVs tend to cluster in labor-intensive and lower value-added industries, restricting the flow of significant positive externalities on DI. In passing note, all control variables in Eq. (4) are consistent with our a priori expectations. For example, whilst the growth in GDP and export volume promotes domestic investment, a higher lending rate deters it.

5.4. Robustness tests

5.4.1. Alternative estimators

In order to ensure the robustness of our results, we re-estimate Eqs. (3) and (4) with the FMOLS, CCR, and DOLS estimators and report the corresponding results in Table 5. Consistent with our ARDL estimates, we find a neutral long-run relationship between FDI and domestic investment across all three estimators, but EJV and CJV, respectively, crowds in and crowds out domestic investment.[10] To put this into context, we estimate that a 1% increase in EJV raises domestic investment by 0.46–0.64%, but a 1% increase in WFFE reduces domestic investment by 0.20–0.31%. The reliability of these estimates is confirmed by the Phillips-Ouliaris τ, Engle-Granger τ, and Hansen instability tests shown in the bottom of Table 5. Meanwhile, the J-B test suggests that the stochastic error term in the FMOLS, CCR, and DOLS estimators in Eqs. (3) and (4) is normally distributed.

In addition, we investigate the possibility that the neutral FDI–domestic investment nexus was brought about by the countervailing effects of different entry modes on FDI. Specifically, we carry out the Wald test for the null hypothesis of H_0: $\delta_1 + \delta_2 + \delta_3 = 0$ and report the relevant test statistic and *p*-value in the bottom of Table 5. In general, we cannot reject the hull hypothesis and conclude that the countervailing effects of different entry modes on FDI tend to offset each other in China.

5.4.2. Individual entry mode

In the preceding discussion, we have included three entry modes in Eq. (4). However, it is informative to also study each entry mode individually and compare those results to our earlier findings. Specifically, we propose the following models:

$$
\begin{aligned}
DI_t &= \rho_1 WFFE_t + X'\beta + \varepsilon_t \\
DI_t &= \rho_2 JV_t + X'\beta + \varepsilon_t \\
DI_t &= \rho_3 EJV_t + X'\beta + \varepsilon_t \\
DI_t &= \rho_4 CJV_t + X'\beta + \varepsilon_t
\end{aligned} \tag{5}
$$

where the estimated coefficient, ρ_i, measures the contribution of entry mode i to domestic investment, and *JV* denotes the sum of EJV and CJV.[11] Since we are holding all things constant in Eq. (5), we can evaluate the effect of entry mode i on domestic investment by comparing the magnitude of ρ_i. Table 6 reports the main results from, and a battery of diagnostic tests on, Eq. (5) using the ARDL and DOLS estimators.[12] In general, we conclude that a cointegrated relationship exists for all specifications in Eq. (5). Importantly, we find that while the coefficient of *JV* is positive, it remains statistically insignificant at all conventional levels. From the outset, this result seems to suggest that joint ventures exert no influence over the FDI–domestic investment nexus in China. However, given the distinct motives and

[10] The CJV–domestic investment nexus appears to be less robust as the coefficient of CJV remains statistically significant only in DOLS, which takes into account both leads and lags.

[11] Given the small share of CJV in our sample, an anonymous referee suggests that it might be worthwhile to also consider the effect of both forms of joint venture on the FDI–domestic investment nexus.

[12] We do not report the FMOLS and CCR estimates as they are quantitatively the same as ARDL and DOLS. Full FMOLS and CCR results are available upon request from the authors.

Table 5
Long-run estimates, by alternative estimator.

	Baseline			Augmented		
	FMOLS	CCR	DOLS	FMOLS	CCR	DOLS
FDI	-0.257	-0.255	-0.109			
	(0.192)	(0.191)	(0.194)			
EJV				0.462**	0.474**	0.644***
				(0.185)	(0.196)	(0.191)
CJV				-0.152	-0.172	-0.334***
				(0.100)	(0.115)	(0.118)
WFFE				-0.311**	-0.299**	-0.204**
				(0.126)	(0.129)	(0.097)
GDP	1.494***	1.493***	1.564***	1.408***	1.410***	1.370***
	(0.102)	(0.089)	(0.139)	(0.092)	(0.081)	(0.138)
EX	0.395***	0.393***	0.269***	0.383**	0.354**	0.167
	(0.093)	(0.090)	(0.097)	(0.158)	(0.166)	(0.160)
RATE	0.017*	0.016	0.008	-0.035**	-0.036**	-0.038**
	(0.010)	(0.010)	(0.014)	(0.016)	(0.017)	(0.019)
Constant	-9.740***	-9.707***	-10.917***	-11.233***	-10.944***	-9.172***
	(1.164)	(1.173)	(1.139)	(1.957)	(2.096)	(2.090)
Adj-R^2	0.996	0.996	0.997	0.996	0.996	0.998
J-B test	0.281	0.361	2.476	0.160	3.426	0.184
PhillipsOuliaris τstatistic	-9.300***	-9.300***	-9.300***	-9.909***	-9.909***	-9.909***
Engle-Granger τ statistic	-10.090***	-10.090***	-10.090***	-11.092***	-11.092***	-11.092***
Hansen instability test	0.503	0.633	0.027	0.999	0.384	0.154
Wald statistic[a]				0.0000	0.0003	0.272
p-vale of the Wald statistic				(0.998)	(0.987)	(0.602)

Note: [a] null hypothesis of the Wald test is $\delta_1 + \delta_2 + \delta_3 = 0$. Standard errors are in parentheses. For DOLS estimates, HAC-adjusted standard errors are reported instead. *,** and *** indicate 10, 5, 1% level of significance, respectively.

Table 6
Individual entry model, by ARDL and DOLS estimators.

	ARDL	DOLS	ARDL	DOLS	ARDL	DOLS	ARDL	DOLS
WFFE	-1.480***	-0.218**						
	(0.527)	(0.110)						
JV			0.623	0.127				
			(0.623)	(0.179)				
CJV					0.749	-0.073		
					(0.997)	(0.115)		
EJV							1.796*	0.446**
							(1.009)	(0.188)
GDP	1.858***	1.578***	1.256***	1.547***	1.484***	1.551***	1.279***	1.514***
	(0.527)	(0.100)	(0.328)	(0.141)	(0.446)	(0.194)	(0.317)	(0.179)
EX	1.096***	0.347***	0.296	0.243**	0.740	0.184	0.171	0.198
	(0.316)	(0.091)	(0.203)	(0.123)	(0.845)	(0.228)	(0.244)	(0.162)
Rate	0.005	-0.002	-0.119	-0.002	-0.135	0.012	-0.215	-0.037*
	(0.019)	(0.015)	(0.089)	(0.017)	(0.152)	(0.011)	(0.138)	(0.022)
Constant	-11.069***	-10.941***	-14.267***	-13.097***	-24.171	-10.213***	-25.800***	-15.386
	(0.659)	(0.410)	(3.972)	(1.825)	(17.862)	(2.115)	(8.841)	(1.725)
ARDL bounds test[a]	9.459***		7.289***		3.227*		4.895***	
Phillips-Ouliaris statistic		-9.419***		-9.149***		-9.388***		-9.091***
Engle-Granger statistic		-10.352***		-9.842***		-10.242***		-9.720***
Hansen instability test		0.033		0.033		0.039		0.040

Note: *,** and *** indicate 10, 5, 1% level of significance, respectively. Standard errors are in parentheses. For DOLS estimates, HAC adjusted standard errors are reported instead. A Each UECM passes key diagnostic tests; we do not report these tests here to conserve space but are available upon request from the authors.

attributes between EJV and CJV discussed in Section 3, these important characteristics might be lost when we treat these two entry modes as a generic group. We avoid such aggregation bias by investigating the potential effect of EJV and CJV individually on domestic investment. Table 6 shows that a 1% increase in EJV raises domestic investment by 0.45–1.80%. In contrast, a 1% increase in WFFE reduces domestic investment by 0.22–1.48%. These findings echo our earlier conclusion that EJVs and WFFEs crowd in and crowd out domestic investment, respectively. In passing note, CJV becomes statistically insignificant at all conventional levels in Eq. (5). However, as highlighted in Section 3, we do not regard this as a problematic result since CJVs only play a minor role in China's vast FDI scene.

5.5. Sub-sample period analysis

In Section 5.1, we identified possible structural breaks in our variables around the time of China's accession to the WTO in 2001. During that period the Chinese economy experienced momentous structural and institutional changes in its FDI regime (Chow, 2003). As a result, foreign participation in the banking, telecommunication, and retailing industries, was allowed for the very first time in China and has been rapidly growing ever since (Zhang, 2014). We capture the potential effect of these changes on the FDI–domestic investment nexus by dividing our sample period into the pre-WTO period (1994Q1–2003Q4) and the post-WTO period (2004Q1–2014Q4). Our selection of this cut-off point reflects partly the CMR unit root

Table 7
The ARDL bounds test and long-run estimates, 1994Q1-2003Q4.

Functional form	Baseline Eq. (3)	Augmented Eq. (4)
	F(DI|FDI;GDP;EX;RATE)	*F(DI|EJV;CJV;WFFE;GDP;EX;RATE)*
Panel I: The bounds test		
Optimal lag structure[a]	(1, 2, 3, 2, 2,)	(1, 1, 0, 3, 3, 0, 2)
F-statistic[b]	6.845***	4.899***
Panel II: Diagnostic tests		
R^2	0.767	0.784
Adj-R^2	0.603	0.591
F-statistics (UECM)	4.693***	4.072***
J-B Normality test	4.217	0.979
B-G LM test	[1] 0.085	[1] 2.418
ARCH LM test	[1] 0.339	[1] 2.406
Ramsay test	[1] 0.894	[1] 0.896
CUSUM	Stable	Stable
CUSUMSQ	Stable	Stable
Panel III: ARDL long-run estimates		
FDI	0.728(0.145)***	
EJV		0.148(0.136)
CJV		-0.094(0.104)
WFFE		0.844(0.214)***
GDP	0.614(0.178)***	0.082(0.286)
EX	0.259(0.108)**	-0.115(0.155)
RATE	-0.068(0.011)***	-0.065(0.013)***
Constant	-6.618(1.501)***	4.672(3.847)

Note: The asterisks ***, ** and * denote the significance at 1%, 5% and 10% levels, respectively. [a] The optimal lag structure of the ARDL models are determined by AIC. [b] The F-statistic is compared to the critical bounds computed by Pesaran et al. (2001) for restricted intercept and no trend. The brackets [ ] is the order of the diagnostic tests. Standard error of ARDL estimates are reported in parentheses. We also reach the same conclusions after comparing the F-test value with the set of critical values provided by Narayan (2005), which is more suitable for a sample size with fewer than 80 observations.

test, and partly, the fact that many of the changes in the FDI regime did not take effect from 2004 and onward.[13]

Consistent with the approach set out for the entire sample period, we provide the ARDL bounds test, a battery of diagnostic tests, and long-run estimates of Eqs. (3) and (4) for the pre- and post-WTO periods. Specifically, Table 7 shows crowding-in between FDI and domestic investment in Eqs. (3) and (4) during the pre-WTO period. In our case, a 1% increase in FDI raises domestic investment by 0.73%, holding all other factors constant. This result is in line with Xu and Wang (2007) and Tang et al. (2008), who also find the crowding-in effect. Braunstein and Epstein (2002) and Huang (2003b) suggest that this effect may stem from the bargaining power of the Chinese government, which persuaded foreign investors to cooperate with state-owned enterprises. Specifically, Eq. (4) indicates that this crowding-in effect is predominantly driven by WFFE rather than EJV, which differs from the finding for the full sample period, where the crowding-in effect was brought about by EJVs.

Turing to the post-WTO period, Table 8 reports a crowding-out effect between FDI and domestic investment in Eqs. (3) and (4). Specifically, a 1% increase in FDI reduces domestic investment by 0.43%, all things being equal. Furthermore, this finding is predominantly driven by WFFEs, coinciding with the Chinese government's

[13] We also detect another structural break around 2008, corresponding to the onset of GFC. However, we are unable to carry out subsample analysis for this subsequent break due to limited observations. Theoretically, it can be argued that since GFC represented a once-off, external shock to the Chinese economy, it exerts only a temporary effect on the FDI–domestic investment nexus.

Table 8
The ARDL bounds test and long-run estimates, 2004Q1-2014Q4.

Functional form	Baseline	Augmented
	Eq. (3)	Eq. (4)
	F(DI|FDI;GDP;EX;RATE)	*F(DI|EJV;CJV;WFFE;GDP;EX;RATE)*
Panel I: The bounds test		
Optimal lag structure[a]	(4, 0, 2, 0, 0)	(1, 2, 0, 1, 2, 0, 0)
F-statistic[b]	4.801***	3.776**
Panel II: Diagnostic tests		
R^2	0.667	0.705
Adj-R^2	0.553	0.582
F-statistic for the UECM	5.822***	5.765***
J-B Normality test	1.256***	2.219
B-G LM test	[1] 0.691	[1] 1.446
ARCH LM test	[1] 0.004	[1] 1.546
Ramsay test	[1] 1.941	[1] 2.167
CUSUM	Stable	Stable
CUSUMSQ	Stable	Stable
Panel III: ARDL long run estimates		
FDI	-0.431(0.117)***	
EJV		0.131(0.103)
CJV		-0.040(0.043)
WFFE		-0.422(0.085)***
GDP	2.162(0.154)***	1.941(0.081)***
EX	-0.608(0.248)**	-0.206(0.111)*
RATE	0.018(0.041)	0.006(0.031)
Constant	-3.184(1.583)*	7.032(1.181)***

Note: The asterisks ***, ** and * denote the significance at 1%, 5% and 10% levels, respectively. [a] The optimal lag structure of the ARDL models are determined by AIC. [b] The F-statistic is compared to the critical bounds computed by Pesaran et al. (2001) for restricted intercept and no trend. The brackets [ ] is the order of the diagnostic tests. Standard error of ARDL estimates are reported in parentheses. We also reach the same conclusions after comparing the F-test value with the set of critical values provided by Narayan (2005), which is more suitable for a sample size with fewer than 80 observations.

Table 9
DOLS estimates, by period.

	1994Q1-2003Q4		2004Q1-2014Q4	
	Baseline	Augmented	Baseline	Augmented
FDI	0.686***		-0.617***	
	(0.096)		(0.119)	
EJV		-0.080		0.018
		(0.286)		(0.161)
CJV		0.155		-0.071
		(0.293)		(0.055)
WFFE		0.645**		-0.506***
		(0.231)		(0.127)
GDP	0.357*	-0.126	2.091***	2.106***
	(0.181)	(0.339)	(0.117)	(0.141)
EX	0.406***	0.429	-0.272*	-0.305
	(0.105)	(0.310)	(0.132)	(0.142)
RATE	-0.062***	-0.045**	0.043	0.038
	(0.006)	(0.018)	(0.044)	(0.041)
Constant	-4.245***	2.781	-4.903***	-5.814***
	(1.370)	(4.824)	(1.193)	(1.749)
Adj-R^2	0.994	0.995	0.998	0.998
J-B Normality test	1.094	0.299	2.712	4.252
Phillips-Ouliaris statistic	-7.326***	-7.217***	-5.140**	-6.002***
Engle-Granger statistic	-4.120	-7.657***	-5.067**	-5.939***
Hansen instability test	0.301	0.435	0.277	0.256

Note: The HAC-adjusted standard errors are reported in the parentheses. *,** and *** indicate 10, 5, 1% level of significance, respectively

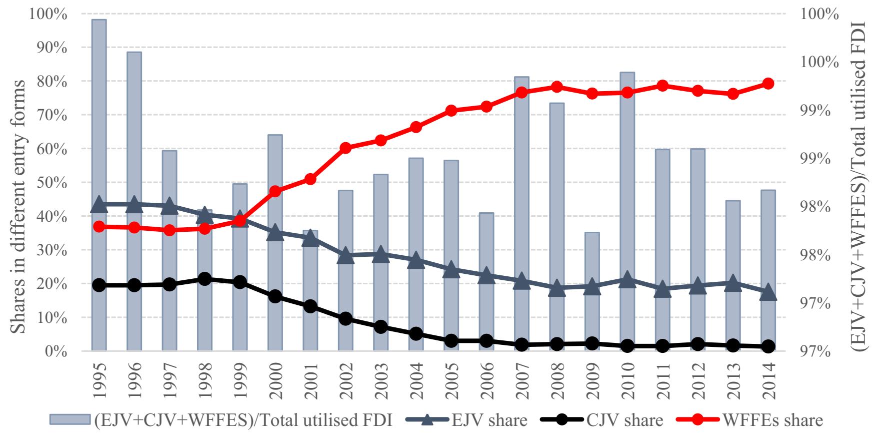

Fig. 1. Utilized FDI, by mode of entry and year, 1995–2014. Source: Authors' calculation based on data from *China Statistical Yearbook*, various issues from 1996 2015.

decision to gradually eliminate institutional barriers to accommodate market-seeking FDI during this period. However, many inefficient Chinese firms were forced out of the market, causing subsequent decline in domestic investment in China.

In general, these results reflect the trends in Fig. 1, where EJV and WFFE were very similar in terms of their FDI proportions between 1995 and 1999. However, WFFE started to rise and EJV falls as early as 2000, by 2004 which is the break point chosen for the sub-sample period, there is already a sizable gap between WFFE (rising in importance) and EJV (declining in importance). From this perspective, it is perhaps not surprising that EJV is not exerting the a priori expected effect on domestic investment. We have attributed this change to either an institutional reason and/or a market reason that took place in China (before 2004) that WFFE as a proportion of total FDI rose 4 years before the change in FDI regime which is supposed to spur higher WFFE.

The robustness of our ARDL results is confirmed by the DOLS estimates of the FDI–domestic investment nexus in both subsample periods. Table 9 reports that the nexus turns from crowding in during the pre-WTO period to crowding out during the post-WTO period. These findings are supported by a battery of diagnostic tests reported at the bottom of Table 9, indicating that our DOLS estimates are largely reliable. In passing note, the DOLS estimates generally exceed their ARDL counterparts, possibly reflecting a richer data generation process under the DOLS estimator.

6. Concluding remarks and policy implications

This paper investigates the fundamental relationship between FDI and domestic investment in China from 1994Q1 to 2014Q4. Unlike its predecessors, this paper is novel in its emphasis on the role of entry mode. Specifically, we find that the neutral nexus in China disappears when entry mode is introduced into the analysis. In fact, we find that EJV crowds in domestic investment, but WFFE crowds it out. These relationships remain valid regardless of the estimation techniques and across different subsample periods. Based on our findings, we show that any study on the FDI– domestic investment nexus in China that ignores entry mode is likely to be incomplete.

In general, we attribute crowding in between EJV and domestic investment in China to positive spillovers brought about by foreign investors benefiting the Chinese firms. Meanwhile, we suspect that crowding out between WFFE and domestic investment in China originated from market-seeking WFFEs displacing their Chinese competitors. These findings carry several policy implications. First, the Chinese government should be selective in granting preferential treatment and fiscal concessions to foreign investors. For example, the Chinese government should discourage the development of WFFEs as it could be detrimental to the growth trajectory of domestic investment. Second, the Chinese government should provide financial and non-financial supports to indigenous firms wanting to form EJVs with foreign investors. As we have shown the promotion of EJVs generates positive spillovers not only for indigenous partners, but also to the wider Chinese economy through various channels. Last, but not the least, our findings echo the recent call by the Chinese government to re-orientate its focus from an export-driven economy to a domestic-led one. However, an important pillar to the success of this attempt hinges on creating an environment that is conducive to sustained growth of domestic investment; and entry mode constitutes a key environmental factor on this front.

Appendix A

See Fig. 1.

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