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Does boardroom gender diversity matter? Evidence from a transitional economy

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ABSTRACT

This research investigates the relationship between board gender diversity and firm financial performance in the context of a transitional economy characterised by an underdeveloped corporate governance system. Using a sample of 120 publicly listed companies in Vietnam covering a 4-year period from 2008 to 2011, we examine this relationship in a dynamic modelling framework, which controls for potential sources of endogeneity. We find that board gender diversity appears to have an effect on firm performance. This finding remains robust when alternative proxies for gender diversity are employed and is consistent with the perspectives of agency theory and resource dependence theory. The number of female directors in the boardroom also matters, supporting the view that if female board representation affects firm outcomes, this effect is more pronounced when the number of female directors increases. It is observed, furthermore, that the marginal positive performance effect of board gender diversity ceases when the percentage of female directors reaches a breakpoint of about 20%. This finding suggests that there is perhaps a potential trade-off between the costs and benefits of board gender diversification. Our findings significantly contribute to the growing literature of non-US based studies, by providing robust empirical evidence from a transitional economy in East Asia.

1. Introduction

Using empirical data from the Vietnamese context, this research contributes to understanding how female representation on boards of directors (BOD) affects a company's financial performance. The topic has become a central focus of corporate governance (CG) rejuvenation efforts around the world, with companies being encouraged to appoint female directors to their boards[1] (Adams & Ferreira, 2009). This raises an important research question as to whether there is a causal relationship between gender diversity on the BOD and firm performance (FP). There has been an increase in the literature on this topic but it relates predominantly to studies in mature markets characterised by well-established CG systems (Adams & Funk, 2012). Several have reported inconclusive results (Campbell & Mínguez-Vera, 2008; Rose, 2007). Moreover, they have not fully addressed potential endogeneity concerns, making inferences about the causal relationship between gender diversity and FP problematic (Terjesen, Sealy, & Singh, 2009). Consequently,

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[1] For example, in 2004 Norway adopted a mandatory gender quota law requiring 40% of positions on the boards of listed companies to be reserved for women (HKEC, 2012). This initiative has motivated many countries in Europe to follow suit, including Belgium (2011), Finland (2005), and Spain (2007). In the Australasian region, Australia (2009), Hong Kong (2012), Malaysia (2011), New Zealand (2012) and others have revised their CG codes to include new "comply or explain" provisions. The new provisions require listed companies to report measurable goals for diversity in their boardrooms, as well as progress in attaining those goals (see e.g., Catalyst, 2012b; HKEC, 2012 for detail).

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the causal effect of board gender diversity on FP, especially in markets characterised by underdeveloped CG systems, remains unclear. The current research, applying a well-structured dynamic modelling approach to control for potential endogeneity concerns, makes an important contribution to understanding how such diversity works in the Vietnamese market and suggests an approach for similar economies.

The issue tends to be more complicated since, as Adams and Ferreira (2009) suggested, the nature of the relationship between board gender diversity and FP is contingent upon whether the firms are well governed. Using a sample of US firms, they contended that because female directors bring tougher monitoring to boardrooms, adding more women directors is likely to provide excessive and unnecessary monitoring for well-governed firms, which may ultimately have a detrimental impact on FP. If so, the subsequent question is whether more gender-diverse boards will improve FP in markets where the companies, which are generally poorly governed, benefit from additional monitoring. Our study addresses this question, contributing to the growing literature of non-US based studies by focusing on Vietnam, a market characterised by an underdeveloped CG system, where the benefits of board diversity may be more pronounced.

It is argued by Carter, Simkins, and Simpson (2003) that the link between board gender diversity and FP is not predicted directly by any single extant theory. Therefore, examining this causal relationship becomes an empirical issue (Carter et al., 2003). As pointed out by Mohan (2014) in a recent comprehensive review paper, however, there are several reasons why we might expect such a causal relationship to exist. Prior research suggests that the presence of women in boardrooms may matter for risk-taking and leadership style, both of which eventually result in effects on FP.[2] If the gender of directors matters for firm outcomes, then female directors may fundamentally differ from their male counterparts in terms of behaviour and personality characteristics.

A recent survey by the United Nations Industrial Development Organization (UNIDO, 2010) confirmed that Vietnamese female entrepreneurs are distinguishable from their male counterparts in regard to both human values and attitudes to risk.[3] These findings are relevant to our current study since Vietnamese female directors are typically appointed from the pool of female entrepreneurs. This being the case, it is plausible that female and male directors in Vietnam will differ in terms of their human values and attitudes to risk. This suggests a causal effect of board female representation on FP in Vietnamese companies. We argue that the UNIDO (2010) findings strengthen the context for the current study and help establish not only the rationale but also the significance of our results for policy implications.

Our study is noteworthy for the following four key reasons. First, it is both relevant and timely due to the board gender diversification policy initiatives currently undertaken by several countries, including the European Union and Australasian countries. Board gender diversity has also emerged as a contemporary policy debate in Vietnam. Accordingly, the significance of the market-based consequences of such policy initiatives is an important policy concern in many countries in the region.

Secondly, our study employs the system generalised method of moments (System GMM) approach, which is considered to be the most appropriate method for exploring the CG–FP relationship in a dynamic framework (Flannery & Hankins, 2013). This estimation technique allows us to control for potential sources of endogeneity which have plagued many earlier studies. Thirdly, the CG–FP relationship and, specifically, the relationship between board gender diversity and FP in the Vietnamese market, are virtually unknown to international scholars.[4] Finally, because the Vietnamese market is characterised simultaneously by a weak CG system and an advanced gender-related institutional context,[5] it provides a unique environment for examining the nexus of board gender diversity and FP.

The rest of this paper is organised as follows. First, we discuss the gender-related institutional environment and the CG context in Vietnam to help readers grasp the significance and background of the research. We then briefly review the relevant literature to develop our main research hypothesis. Data, data sources, and method are described next. Finally, we present the empirical results and conclude the paper with a discussion of the contributions and limitations of the study.

2. Background

Corporate governance is a new concept for Vietnam and there is also no equivalent Vietnamese terminology that fully explains CG. The term 'corporate governance' is translated as 'quản-trị-công-ty', similar in meaning to 'company administration' (OECD, 2006). The CG system in Vietnam is in the initial stages of development (World Bank, 2006a) and the current situation can be characterised as

[2] For example, Adams and Funk (2012) documented that female and male directors are systematically different in their core values and attitudes to risk. The subsequent question is how financial markets evaluate these differences. Adams et al. (2011) found that market reaction to the appointment of female directors is—on average—significantly positive, and consistently greater than it is to the appointment of their male counterparts. Mohan and Chen (2004), however, documented that the initial public offering (IPO) markets do not appear to distinguish between female- and male-led IPOs when evaluating them.

[3] For instance, while Vietnamese male entrepreneurs are risk-taking investors and tend to make decisions by themselves, their female counterparts—due to cultural tradition and their social role—tend to consult their family members on important business decisions (UNIDO, 2010). Furthermore, the perseverance and determination to succeed of Vietnamese female entrepreneurs appear to be greater than those of their male counterparts. As goal-oriented entrepreneurs, Vietnamese women also take their businesses seriously, participate in entrepreneurial organisations, and readily grasp how to use informal means to promote their own businesses (UNIDO, 2010).

[4] The latest review paper on the theme of CG in emerging markets conducted by Claessens and Yurtoglu (2013) does not include any information about Vietnam. Another recent meta-analysis paper concerning corporate board–FP relationship in the Asian region conducted by Van Essen et al. (2012) similarly provides no information about Vietnam. In the most recent comprehensive review paper by Terjesen et al. (2009) dealing with the topic of female directors in the boardroom, there is no research using Vietnamese company data among more than 400 relevant publications. We also conducted a simple 'survey' at the end of 2012 to look for publications of the CG and FP relationship in Vietnam. We followed Love (2011) and used the key words 'corporate governance' + 'performance' + 'Vietnam' to search www.GoogleScholar.com, www.SSRN.com, and the Proquest5000 database. Generally speaking, the search results showed that there was no empirical research considering the case of Vietnam.

[5] See Background section for more detail.

follows: (i) CG regulations are less developed and enforced (World Bank, 2006a); (ii) public awareness regarding CG is poor (Freeman & Nguyen, 2006); (iii) the role of the state sector is predominant; (iv) the protection of private property rights is weak; and (v) both internal and external governance mechanisms are limited (Le & Walker, 2008; Nguyen, 2008; World Bank, 2006a).

To establish an effective CG system and improve public awareness regarding CG, the Vietnamese government has adopted the legal principles of Anglo-American jurisdictions in order to establish a regulatory system for the CG practices of Vietnamese companies (Le & Walker, 2008). Specifically, the Law on Enterprises was enacted in 2005 and became officially binding as of July 2006 (hereafter the LOE 2005), marking a turning point in the development of business freedom and the legal framework of CG practices in Vietnam (Bui & Nunoi, 2008).

Under the LOE 2005, the Vietnamese Ministry of Finance promulgated the Code of Corporate Governance for Listed Companies (the Code) in March 2007, updated in July 2012, reflecting most of the OECD Principles of CG (the OECD Principles).[6] However, it is noteworthy that while the OECD Principles are a flexible, principle-based approach to governance (also referred to as the 'comply or explain' approach), the Code is mandatory for all publicly listed companies in Vietnam (Le & Walker, 2008). The Code also provides guidance on board diversity in terms of professional background, management experience, and industry knowledge. Nevertheless, it is interesting to note that the Code does not make any reference to either gender or racial diversity.

In complying with the LOE 2005 (National Assembly, 2005) and the Code (MOF, 2007), the typical governance structure of a Vietnamese listed company follows a two-tier model and consists of four governance bodies: (i) a general meeting of shareholders (GMS); (ii) a board of directors (BOD); (iii) a chief executive officer (CEO); and (iv) a board of supervisors (BOS). The GMS, the most powerful body of a publicly listed company, establishes the company's constitution and elects the members of both the BOD and BOS. In accordance with the company's constitution, the BOD chairperson may be elected by either BOD members or the GMS. As stipulated by the LOE 2005, the BOD—consisting of three to eleven members—is responsible for guiding and establishing the company's business strategies as well as monitoring managerial decisions. Specifically, the LOE 2005 clearly stipulates four major duties of the BOD, including: (i) making decisions regarding management strategies; (ii) nominating the CEO and approving senior executive positions; (iii) monitoring daily managerial operations; and (iv) proposing matters for the consideration of the GMS. In comparison with the German internal CG model, the BOD of Vietnamese companies has a more direct role in monitoring daily management (Le & Walker, 2008).

The LOE 2005 provides that a BOS must be established in companies which have more than eleven individual shareholders or at least one institutional shareholder holding more than 50% of the company's equity. The membership of a BOS must range from three to five members who need not be shareholders or employees of the company. Unlike the one-tier board structure in Anglo-American jurisdictions where a supervisory committee is composed and nominated by the BOD, the members of a Vietnamese BOS are elected by the GMS and function independently from the BOD (Bui & Nunoi, 2008). According to the LOE 2005, more than half of the BOS's membership must reside permanently in Vietnam and at least one member must be an accountant or auditor.

The major role of the BOS is to make an internal assessment of the annual financial statements and supervise the performance of both BOD and CEO. However, the LOE 2005 does not stipulate what specific form of supervision is required and how the BOS should implement its decisions (Bui & Nunoi, 2008). The absence of clear legal guidance for the BOS on what and how to supervise the BOD has had the result that the BOS's supervisory role in Vietnamese companies is ineffective (World Bank, 2006a). In consequence, the BOS in Vietnamese companies, in reality, appears to exist in form rather than in substance (Bui & Nunoi, 2008).

Despite the efforts made by the government to improve the standard of governance practised by publicly listed companies, the CG system in Vietnam still remains underdeveloped. Indeed, Vietnam is ranked 166th out of 183 economies for the strength of investor protection (World Bank, 2012). The most recent CG scorecard for 2011, conducted by the IFC (2012), reported that the average CG score in Vietnam is only 42.5%, which is much less than those of other markets across the Asian region.[7]

Regarding the gender-related institutional environment in Vietnam, UNIDO (2010) argues that the country has been strongly influenced by Confucian gender ideologies in which women are subordinated to men. However, Vietnamese companies nowadays enjoy an advanced gender-related institutional environment in which women's rights and gender equality are constantly promoted. As a Marxist–Leninist one-party state, Vietnam has pursued 'a socialist-oriented market economy' in which the state sector rather than market forces plays the decisive role in controlling the economy. In such an economic structure, the government intervenes strongly and directly in the economy in order to achieve the socialist ideals of citizens' equality and, to a lesser extent, gender equality.

Gender equality, therefore, is considered to be one of the central goals of this communist state's socio-economic development strategies (Knodel, Vu, Jayakody, & Vu, 2004). Since 1945, the Vietnam Communist Party has been strongly committed to achieving this goal by adopting gender-based interventions. In 2002, for example, the Vietnamese government proclaimed a *National Strategy for the Advancement of Women to 2010* that identifies high priorities for achieving equal rights for women in labour, employment, education, health, and economic participation (Asian Development Bank, 2005). The National Strategy on Gender Equality 2011–2020, adopted in 2010, also specifies objectives for the participation of women in leadership and management (World Bank, 2011).

Through concerted efforts for gender equality, Vietnam has achieved key gender equality indicators extremely well in comparison with other East Asian countries at a similar, or even higher, level of GDP per capita (World Bank, 2011). For instance, the World Bank (2006b) assessed Vietnam as one of the countries in the world that had achieved the highest rate of economic participation by women and the highest participation of women in state power structures, such as parliament, in the East-Asian region. More recently, the

[6] The OECD Principles of CG were approved by Organisation for Economic Co-operation and Development Ministers in 2004 and have since become an international benchmark of CG practices for policy makers worldwide.

[7] For example, the average scores of Thailand (in 2011), Hong Kong (in 2009), and the Philippines (in 2008) are 77%, 73% and 72%, respectively (IFC, 2012).

World Bank (2011) reported that the participation rate of Vietnamese women in the labour force ranked among the highest for countries in the region and that the gender gap in earnings was lower in Vietnam than in many other East Asian countries.

Vietnam has also made considerable progress in reducing gender-related hindrances in the business environment for female entrepreneurs (UNIDO, 2010). More specifically, UNIDO (2010, p. 12) reported that this organisation "did not find any significant difference in perceived gender-based bias of male and female entrepreneurs in getting collateral, entering networks, acquiring new contracts, employing workers and dealing with authorities." This situation may facilitate economic participation and promotion opportunities for Vietnamese women, which in turn may help to extend the pool of qualified women from which the most suitable candidates for director will be chosen.

According to Grosvold and Brammer (2011) and Terjesen and Singh (2008), national institutional systems, such as the socio-economic and political structure, legal background, governance system, and cultural foundation, among others, constitute important antecedents for female representation in boardrooms as well as opportunities for women to advance in their careers. For this reason, Grosvold and Brammer (2011) recommend that the national institutional environment should be completely incorporated in studies on boardroom diversity. The institutional environment in Vietnam, on the one hand, is remarkable for its underdeveloped CG system and on the other hand, is characterised by advanced gender-related institutions. Together, these distinctive institutional features make Vietnam an interesting case to study.

3. Literature review and hypothesis development

Theoretically, the link between board gender diversity and FP is not predicted directly by any single theory, including agency theory and resource dependence theory[8] (Carter, D'Souza, Simkins, & Simpson, 2010). However, both these theories do provide insight into the link and imply the possibility that board gender diversity affects firm value (Carter et al., 2010). In fact, there is a small but developing literature documenting that female board representation matters for firm outcomes (Adams & Funk, 2012).

According to agency theory, the monitoring function of the BOD plays an extremely important role in mitigating principal–agent conflicts, which ultimately affect FP (Fama & Jensen, 1983; Jensen & Meckling, 1976). Recent empirical studies suggest that greater gender diversity on boards has the potential to strengthen this monitoring function. For example, Adams and Ferreira (2009) and Adams, Nowland, and Grey (2011) reported that female directors tend to have better monitoring ability because they are able to think independently and are not affected by the so-called old-boys' club syndrome. Greater gender diversity on boards may also provide better monitoring since female director representation helps to improve managerial accountability, such as improving board meeting attendance and CEO responsibility (Adams & Ferreira, 2009). As a result, female directors may act as additional independent directors who help to improve the monitoring function of the BOD (Adams & Ferreira, 2009).

However, it is worth noting that even if boards with more gender diversity do improve the monitoring function of the BOD, it does not necessarily follow that this improvement will result in better FP. This is because the potential effect of gender diversity on FP is contingent upon the quality of firm governance. Adams and Ferreira (2009) suggested that weakly governed companies may benefit from including more women on their boards, enhancing additional monitoring and improving firm value. In support, Gul, Srinidhi, & Ng (2011, p. 314) argue that greater gender diversity on boards acts as a "substitute mechanism for corporate governance that would be otherwise weak", and this in turn may lead to improved performance. Conversely, board gender diversity seems to have a harmful effect on the FP of well-governed firms because of unnecessary, excessive monitoring (Adams & Ferreira, 2009).

Resource dependence theory suggests that the security of firms' vital resources as well as the linkage between firms and their external environment can be improved by an increase in the size and diversity of the BOD (Goodstein, Gautam, & Boeker, 1994; Pfeffer, 1973). In other words, firms with larger and/or more diverse boards may have advantages in obtaining and maintaining their important resources, including: (i) the human capital of board members (knowledge, skills, and talent); (ii) advice and counsel; (iii) channels of communication; and (iv) legitimacy (Hillman & Dalziel, 2003; Pfeffer & Salancik, 2003). In fact, it is documented in the CG literature that more gender-diverse boards may help to extend these firms' vital resources (Liu, Wei, & Xie, 2014). Hillman, Cannella, and Harris (2002) argued that diversifying the BOD by adding more women would help companies to gain legitimacy as gender equality becomes increasingly one of the widely accepted social norms.

In a similar vein, female directors may broaden the human capital and channels of communication of the BOD by offering additional insight into firms' strategic issues, especially those that relate to female employees, consumers, and business partners (Daily, Certo, & Dalton, 1999). It follows that female representation in boardrooms should improve information processing, leading to higher quality decisions and ultimately better FP (Dezsö & Ross, 2012; Rose, 2007). However, greater boardroom gender diversity may not necessarily result in more effective boards (Carter et al., 2003). More specifically, greater board gender diversity may lead to several difficulties in reaching a consensus on strategy decisions and in implementing monitoring functions effectively, since the greater the diversity of the BOD, the greater the potential that conflict of interests may occur (Goodstein et al., 1994).

In summary, although both theories suggest that the link between board gender diversity and FP appears to be a real possibility (Carter et al., 2010), the nature of the link remains unclear (Carter et al., 2010; Erhardt et al., 2003; Rose, 2007). The empirical question that needs to be answered is, if the link between board gender diversity and FP does exist, does female director representation make the difference? Prior empirical studies on this topic, predominantly conducted in developed markets, provide us with inconclusive evidence (Campbell & Mínguez-Vera, 2008; Mohan, 2014; Rose, 2007).

[8] Therefore, "until a theoretical framework that predicts the nature of the relationship is developed", examining the board gender diversity–FP nexus is an empirical issue (Carter et al., 2003, p. 38). Nevertheless, among several theories from various fields, resource dependence theory provides "the most convincing theoretical arguments for a business case for board diversity" (Carter et al., 2010, p. 398).

Some argue that the relationship between gender diversity and performance is positive (Campbell & Mínguez-Vera, 2008; Carter et al., 2003; Dezsö & Ross, 2012; Erhardt, Werbel, & Shrader, 2003), or negative (Adams & Ferreira, 2009; Ahern & Dittmar, 2012), while others see evidence of no significant relationship at all (Carter et al., 2010; Rose, 2007). We argue that such mixed empirical evidence reflects the differences in research contexts and econometric techniques used. For instance, given that women tend to work for better performing companies (Farrell & Hersch, 2005), studies that link gender diversity to FP should treat gender diversity as an endogenous variable (Adams & Ferreira, 2009; Carter et al., 2010; Dezsö & Ross, 2012). This implies that ignoring the endogenous nature of the gender diversity–FP connection makes empirical estimations problematic.

Given that the extant theoretical framework and prior empirical findings do not suggest a clear outcome for the board gender diversity–performance nexus, our theory-based analysis will be adjusted by the Vietnamese CG context. Accordingly, if the performance effect of greater gender diversity on boards appears to be more pronounced in firms with weak governance (Adams & Ferreira, 2009; Gul et al., 2011), we can plausibly infer that Vietnamese firms, characterised by underdeveloped governance practices, may greatly benefit from adding female directors to their boards. In other words, we argue that if female directors provide greater monitoring expertise, which is more valuable in a weak CG environment (Adams & Ferreira, 2009; Adams et al., 2011; Gul et al., 2011), it may be expected that Vietnamese listed companies with more gender-diverse boards will enjoy better financial performance. Therefore, we propose our main research hypothesis as follows:

Hypothesis. Board gender diversity has a significantly positive effect on the financial performance of Vietnamese listed companies.

4. Data and method

4.1. Sample and data sources

The Ho-Chi-Minh Stock Exchange (HOSE) and the Hanoi Stock Exchange (HNX) are two stock markets in southern and northern Vietnam, respectively. Of 275 non-financial companies listed on these two bourses at the end of 2008, 122 companies have relatively complete data on key variables during the 4-year period from 2008 to 2011.[9] Hence, a panel dataset comprising 488 firm-year observations is used as the initial dataset. However, to ensure that our findings are not driven by the outliers of Tobin's Q, we drop firm-year observations within the first and beyond the 99th percentiles, as suggested by Balatbat et al. (2004), Giroud and Mueller (2010), Kuo et al. (2014), among many others. Consequently, our final sample includes 479 firm-year observations. Company financial data are sourced from Thomson One Banker (*Worldscope database*). Block-holder ownership data are extracted from *Thomson One Banker* (*Ownership module*). We also collate block-holder ownership information from the companies' annual reports. Data for board structures are hand-collected from the firms' annual reports downloaded from *FPT-Ez-search Online Information Gateway* and *Vietstock*. The list of publicly listed companies, classified according to the Industry Classification Benchmark (ICB), is provided by *StoxPlus Corporation*.

4.2. Variables

4.2.1. Dependent variable

Following prior studies (e.g., Mohan & Ruggiero, 2007; Reddy et al., 2008), we employ the Tobin's Q ratio as the FP measure. As a market-based measure of financial performance, a Tobin's Q ratio higher than one reflects investors' expectation that the company has powerful comparative advantages or good growth opportunities (Campbell & Mínguez-Vera, 2008; Rose, 2007). In contrast, a Tobin's Q ratio smaller than one indicates poor utilisation of company resources (Campbell & Mínguez-Vera, 2008; Rose, 2007). We follow prior studies (Chen et al., 2006; Nguyen et al., 2014) and compute the Tobin's Q ratio as the sum of the market value of a firm's stock and the book value of debt divided by the book value of total assets. To mitigate the potential effects of outliers, we transform Tobin's Q into natural logarithmic form.

4.2.2. Explanatory and control variables

Our independent variable of interest is gender diversity. We first use the percentage of female directors on boards (*female*) as a proxy for gender diversity. In order to check the robustness of the estimations, we follow Campbell and Mínguez-Vera (2008) and employ two alternative proxies for gender diversity: (i) a gender diversity dummy variable (*d1women*); and (ii) the Blau index for gender (*blau*). The variable *d1women*, which distinguishes companies with at least one female director on their boards from those without, allows us to answer the question whether the presence of women in boardrooms in itself has an impact on FP.

Meanwhile, the variable *female* enables us to examine the effect that board gender balance has on FP. The *Blau index* for gender combines both of the above aspects of diversity, including the variety (measured by *d1women*) and the balance (measured by *female*)

[9] Consistent with the previous literature, financial firms and banks are excluded from the sample because they function under strict regulations that may have various influences on their CG mechanisms. It should be noted that the number of non-financial listed companies in Vietnam has increased by 317 companies over a 4-year period, from 275 in 2008 to 592 in 2011. Because we examine a 4 consecutive year dataset from 2008 to 2011, these 317 companies listed after 2008 are not included in our sample, leaving an available population of 275 companies listed in 2008. The reason for considering data from 2008 is that these Vietnamese listed companies, in compliance with the Code, have consistently provided CG information in their annual reports from fiscal year 2008 onward. In relation to the whole population of 275 non-financial companies listed in 2008, the beginning year in the 4 consecutive year dataset, our final sample accounts for approximately 44% and is the largest possible dataset we can obtain.

(Campbell & Mínguez-Vera, 2008). Following Blau (1977, as cited in Harrison and Klein, 2007), we also calculate the Blau index for gender as $(1-\sum_{i=1}^{2}P_i^2)$, where $i = (1, 2)$ is the number of gender categories (two); P_i is the proportion of board members in each category. The minimum and maximum values of the Blau index for gender are zero (perfectly homogeneous boards) and 0.5 (perfectly heterogeneous boards), respectively.

Moreover, in order to capture the potential effect of the number of female directors, we follow Liu et al. (2014) and employ a dummy variable (denoted as *d2women*) that takes a value of one if there are at least two female directors and zero otherwise. We also control for other board and firm characteristics that may have effects on FP, such as the percentage of non-executive directors on board, duality, board size, and block-holder ownership. In line with prior studies, we treat firm size, leverage (e.g., Chen et al., 2003), firm age (e.g., Reddy, Locke, Scrimgeour, & Gunasekarage, 2008), industry dummies, and year dummies as control variables. In addition, we use the 1-year lagged dependent variable, recommended by Wintoki et al. (2012) and others, as an explanatory variable to control for the dynamic nature of the CG–FP relationship. The detailed definitions and acronyms of the variables used in this study are summarised in Table 1.

4.3. Method

4.3.1. Model specification

As suggested by Wintoki, Linck, and Netter (2012), the model specification for estimating the CG–FP relationship in a dynamic framework[10] is described as follows:

$$y_{it} = \alpha + \sum_{s=1}^{k} \varphi_s y_{it-s} + \beta X_{it} + \delta Z_{it} + year\ dummies + industry\ dummies + \eta_i + \varepsilon_{it} \tag{1}$$

where: i indexes observational firms and t indexes time; φ, β, and δ are vectors of coefficients on lagged dependent variables (y_{it-s}), CG variables (X_{it}) and control variables (Z_{it}), respectively; η_i represents unobserved time-invariant firm effects; ε_{it} is a random error term; and k is the number of dependent variable lags. Pham et al. (2011) and Wintoki et al. (2012) suggested that two lags of the dependent variable ($k = 2$) are sufficient to capture all information from the past.

In order to check this argument empirically, we follow Wintoki et al. (2012) and run an OLS regression of current performance on two lags of past performance (i.e., y_{it-1} and y_{it-2}), controlling for X_{it} and Z_{it}. We find that the coefficient on y_{it-2} is not statistically significant at the 5% level, suggesting that one lag is enough to capture the dynamic nature of the CG–FP relationship. Thus, the 1-year lag of Tobin's Q will be employed as an explanatory variable in our official regression models. This is in line with the studies of Adams and Ferreira (2009); Dezsö and Ross (2012); and Nguyen, Locke, and Reddy (2014). When $k = 1$, Eq. (1) can be specifically written as follows:

$$\begin{aligned} lnq_{it} = {} & \alpha + \varphi lnq_{it-1} + \beta_1 female_{it} + \beta_2 nonexe_{it} + \beta_3 dual_{it} + \beta_4 lnbsize_{it} + \beta_5 block_{it} \\ & + \delta_1 lnfage_{it} + \delta_2 fsize_{it} + \delta_3 lev_{it} + year\,dummies + industry\,dummies + \eta_i + \varepsilon_{it}. \end{aligned} \tag{2}$$

Eq. (2) is the base-line model in our study. In order to compare with previous studies and emphasise the dynamic nature of the CG–FP association, we estimate Eq. (2) by using alternative estimation models, including: (i) a static OLS model; (ii) a fixed-effects model; (iii) a dynamic OLS model; and (iv) a two-step System GMM model. In these models, φ, β_i ($i = 1, 2, 3, 4, 5$), and δ_j ($j = 1, 2, 3$) are estimated coefficients of explanatory variables and control variables, respectively. For the static OLS model and fixed-effects model, it is assumed that $\varphi = 0$. It should be noted that the industry dummies, by construction, are not included in estimations with fixed effects, including models (ii) and (iv). In an unreported additional analysis, we find that all coefficients on industry dummy variables in model (iii) are statistically insignificant at the 5% level. The null hypothesis of the Wald test that the coefficients on all industry dummies are simultaneously equal to zero cannot be rejected at any conventional levels of significance. For this reason, we exclude industry dummies from the official OLS estimations and rerun models (i) and (iii) without industry dummies. Following Schultz, Tan, and Walsh (2010); and Wintoki et al. (2012), we assume that firm age (*lnfage*) and *year dummies* are exogenous.

4.3.2. Estimation approach

Most prior studies on the CG–FP nexus commonly employed the fixed-effects (FE) approach and/or the traditional instrumental variable (IV) approach to mitigate potential endogeneity concerns arising from unobserved time-invariant heterogeneity and/or simultaneity. However, these techniques are not designed to deal with dynamic endogeneity, which very likely arises in the board structure to performance relationship in general (Wintoki et al., 2012) and in the gender diversity to performance relationship in particular (Adams & Ferreira, 2009; Dezsö & Ross, 2012). In addition, applying the traditional IV approach, which requires identifying reliable external instruments, is no easy task (Flannery & Hankins, 2013). It is therefore extremely difficult, if not impossible, to look for a

[10] The theoretical arguments of Harris and Raviv (2008), Hermalin and Weisbach (1998), and Raheja (2005) imply that the relationship between CG and FP is dynamic in nature; that is, current CG characteristics and performance are affected by firms' past performance. This is confirmed by Schultz et al. (2010) and Wintoki et al. (2012), who suggest that the appropriate empirical model for the CG–FP association should be a dynamic one, in which lagged performance is used as one of the explanatory variables, rather than the static model applied by many prior studies. In the context of CG literature, this dynamic approach has been recently applied in studies of the board structure–performance nexus (e.g., Wintoki et al., 2012), the determinants of board structure (e.g., Chen, 2014) or the CG–FP nexus (e.g., Nguyen et al., 2014).

Table 1
Variables and definitions.

Variables	Acronym	Definition
Dependent variable		
Tobin's Q ratio	*lnq*	Tobin's Q is the sum of the market value of firm's stock and the book value of debt divided by the book value of its total assets. The natural logarithm of Tobin's Q (*lnq*) is used as the regressant.
Gender diversity (variable of interest)		
Percentage of female directors (%)	*female*	The ratio of female directors to total number of directors.
Blau index for gender	*blau*	Blau index for gender $= 1 - \sum_{i=1}^{2} P_i^2$, where $i = (1, 2)$ is the number of gender categories (two), P_i is the proportion of board members in each category.
Dummy variable for gender diversity (1)	*d1women*	Dummy variable that takes a value of one if there is at least one female director, and zero otherwise.
Dummy variable for gender diversity (2)	*d2women*	Dummy variable that takes a value of one if there are at least two female directors, and zero otherwise
Board structure variables		
Percentage of non-executive directors (%)	*nonexe*	The ratio of non-executive directors to total number of directors.
Duality	*dual*	Dummy variable that takes a value of one if the chairperson is also the CEO, and zero otherwise.
Board size	*lnbsize*	Board size is the total number of directors. The natural logarithm of board size (*lnbsize*) is used in our models.
Other control variables		
Block-holder ownership (%)	*block*	The ratio of ordinary shares held by shareholders who own at least 5% of the total number of a company's common stocks.
Firm age	*lnfage*	The natural logarithm of the number of years from the time the company first appears on the HOSE or the HNX.
Firm size	*fsize*	The natural logarithm of the book value of total assets.
Leverage (%)	*lev*	The ratio of the company's total debt to its total assets.
Lagged dependent variable	*laglnq*	One year lagged Tobin's Q ratio (in natural logarithmic form).
GDP growth	*gdpgrow*	GDP growth is measured by the annual percentage growth rate of GDP calculated at market prices.
Year dummy variables	*year*	A dummy variable for each year from 2008 to 2011. One year dummy is treated as the benchmark category to avoid dummy variable trap.
Industry dummy variables	*industry*	A dummy variable for each of the eight industries, including: Basic Materials; Consumer Goods; Consumer Services; Health Care; Industrials; Oil & Gas; Technology; and Utilities. One industry dummy is treated as the benchmark category to avoid dummy variable trap.

set of multiple external instruments for the current study in which almost all explanatory variables are considered to be endogenously determined.

Given the unavailability of appropriate external instruments for CG research, the two-step System GMM estimator—proposed by Blundell and Bond (1998)—constitutes the most feasible solution for dealing with endogeneity issues arising from a dynamic panel setting (Antoniou et al., 2008; Nakano & Nguyen, 2012). This technique, on the one hand, allows us to employ internal instruments available within the panel itself (Blundell & Bond, 1998), facilitating our empirical estimation process. On the other hand, it allows us to cope with "the combination of a short panel, a dynamic dependent variable, fixed effects and a lack of good external instruments" (Roodman, 2009b, p. 156). Indeed, simulation analyses recently undertaken by Flannery and Hankins (2013); and Zhou et al. (2014) documented that the System GMM emerges as the best-performing estimator across common data features encountered in our dataset, including: (i) short panel; (ii) endogenous explanatory variables; and (iii) dynamic panel bias. More importantly, by construction the System GMM estimator allows for mitigating the problem of the slow-changing characteristics of independent variables, which renders the FE estimator powerless (Antoniou, Guney, & Paudyal, 2008).

5. Empirical results and discussion

5.1. Descriptive statistics

Table 2 reports the descriptive statistics of the key dependent and independent variables used. Tobin's Q values range from 0.20 to 2.96, with an average value of 0.85. The median Tobin's Q of 0.78 means that for half the observations, the Tobin's Q is less than or equal to 0.78. Furthermore, the median Tobin's Q of 0.78 is very close to the mean Tobin's Q of 0.85, both of which are less than one. This suggests that in terms of central tendency, the market value of the listed companies during the sampling period is lower than the book value. On the one hand, this result may reflect the negative expectation of investors in response to the ineffective use of scarce company resources. On the other hand, it may also reflect the variations of the Vietnamese stock exchange during the crisis period of 2008–2011.

Appendix 1 demonstrates that the mean and median values of Tobin's Q closely follow the fluctuations of the Vietnamese Stock Index (*VNIndex*) across the years from 2008 to 2011. It is evident from Appendix 1 that the mean and median of Tobin's Q—on a year-by-year basis—are also smaller than one when the *VNIndex* annual growth rates are negative. Consequently, we believe that the negative trend of the market during this crisis period is a possible explanation why Vietnamese listed companies were undervalued by investors.

Table 2
Descriptive statistics of the trimmed sample.

	Obs	Mean	Median	Sd	Min	Max
Tobin's Q ratio	479	0.85	0.78	0.39	0.20	2.96
Female directors (%)	472	12.06	9.09	13.76	0.00	66.67
Dummy variable for gender diversity	472	0.51	1.00	0.50	0.00	1.00
Non-executive directors (%)	479	48.91	42.86	20.76	0.00	100.00
CEO duality	479	0.32	0.00	0.47	0.00	1.00
Board size	479	5.81	5.00	1.29	4.00	11.00
Block ownership (%)	478	43.92	49.28	20.86	0.00	86.89
Firm age (year)	479	3.34	3.00	2.04	0.00	11.00
Firm size [Ln(Total assets)]	479	27.24	27.22	1.20	24.11	30.55
Leverage (%)	479	29.22	28.00	20.27	0.00	75.69

Note: For descriptive purposes, *Firm age* and *Board size* are reported in levels. The variables are as defined in Table 1.

The mean percentage of female directors is 12.06%, which is twice as many as that reported by Sussmuth-Dyckerhoff et al. (2012) for the Asian region (6%). Furthermore, as reported by Catalyst (2012a), the mean percentage of female directors in Vietnam is far larger than that of other countries in the region, such as China (8.50%), Hong Kong (9%), Indonesia (4.50%), Japan (0.90%), Malaysia (7.80%), Singapore (6.90%), South Korea (1.90%), and Thailand (8.70%). Given that the institutional environment has an important influence on the social role and boardroom representation of women (Grosvold & Brammer, 2011), the higher ratio of female directors in Vietnamese companies appears to be the direct outcome of a better, more gender-diverse institutional context, mentioned earlier in Section 2.

The mean of the variable *d1women* is approximately 0.51 suggesting that 51% of companies in the sample (equivalent to 239 out of 472 observations, as reported in Table 3) have at least one female director on their board. Arguably, this proportion is much higher than that reported by Campbell and Mínguez-Vera (2008) for Spain (23.70%) and by CGIO. (2011) for Singapore (40%). This result is a reflection of the high proportion of women in the labour force in Vietnam (World Bank, 2011), which may contribute to higher gender diversity in the boardroom than would otherwise be the case.

Nevertheless, female representation in Vietnamese boardrooms is still low. As reported in Table 3, of 239 cases with at least one female director, only 75 (≈31%) have two or more women on the board. The number of cases with at least three female directors is negligible (20 cases). Table 3 also provides detailed information about the frequency of female directors by board size. It can be observed from Table 3 that cases with one or two female directors on the board tend to be those which have a board membership ranging from five to seven.

On an average basis, non-executive directors account for about 49% of total directors, similar to the ratio found by the IFC (2011) for the Vietnamese market. The minimum percentage of non-executives is zero, although to ensure board independence (MOF, 2007), the Code requires that at least one-third of Vietnamese listed company directors must be non-executives. This situation reflects the fact that while one or more companies in the sample have failed to comply with the minimum level of non-executives on the BOD, the others have achieved well above the threshold. Among companies in the sample, only 32% of the board chairpersons are also the CEOs, indicating that dual roles are less common in Vietnam. This ratio is in agreement with that reported by the IFC (2011) in its survey of the Vietnamese market.

The average number of board directors is approximately six, similar to the ratio reported by the IFC (2011) and, as reported by The Korn/Ferry Institute (2012), much smaller than the average board size of other countries in the Asian region, such as China

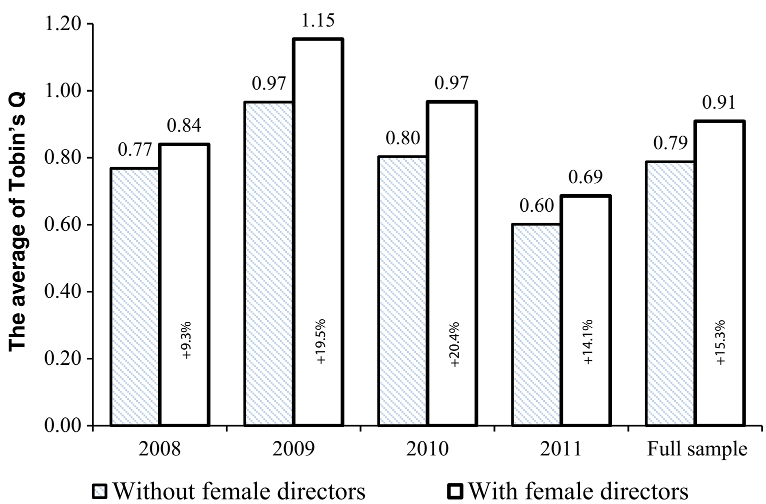

Fig. 1. The average values of Tobin's Q of companies with and without female directors.

Table 3
The frequency of female directors by board size.

Board size (persons)	Number of female directors in boardroom (persons)					Total
	0	1	2	3	6	
4	1	5	0	0	0	6
5	163	111	19	0	0	293
6	14	8	17	5	0	44
7	43	25	15	6	0	89
8	3	3	1	5	0	12
9	8	8	1	0	1	18
10	1	0	2	3	0	6
11	0	4	0	0	0	4
Total	233	164	55	19	1	472

Note: *Board size* is as defined in Table 1.

(11.60), Hong Kong (11.50), India (10.80), Malaysia (9.00), and Singapore (8.60). Table 3 shows that there are from five to seven members on the boards of most companies in our sample. It can be seen that the board size of companies in Vietnam is: (i) in compliance with the requirement of the Code that the boards should have from three to eleven members; and (ii) within the optimal threshold of board size—no more than eight members—recommended by Jensen (1993) for board effectiveness.

The mean value of firm age (the period of time from the initial public offering) is 3.34, reflecting the fact that listed companies in the sample are very young. This can partly explain their lack of experience in dealing with CG issues pointed out in recent IFC reports (IFC, 2011, 2012). Notably, about 44% is the average percentage of ordinary shares held by shareholders who own at least 5% of the total number of a company's common stocks. Although this number is lower than that of Singapore (60%) and Malaysia (47%) (see Mak & Kusnadi, 2005), it is still much higher than that of western developed markets, such as the US and UK markets. This finding is in line with Chen and Huang (2014), who documented that many emerging markets are characterised by highly concentrated ownership structures. While this may indicate that the ownership structure of listed companies in Vietnam is highly concentrated, it should be noted that this ratio varies considerably from zero to approximately 87%.

Table 4 reports that the variation in the ratio of female directors within firms (5.34%) is lower than that across firms (12.85%). This result suggests that the percentage of female directors does not vary greatly over time. In other words, *female* is a slow-changing variable, given that if a variable does not vary across time, the within-standard deviation will be zero. The remaining CG variables, including *nonexe*, *dual*, *lnbsize*, and *block*, share a similar characteristic. These findings are consistent with Brown et al. (2011), who reported that most CG variables do not change over time, which leads to a reduction in the statistical power of CG research. Taking into account the slow-changing feature of these variables, we employ the two-step System GMM as the main approach. The reason for this is that the two-step System GMM, by construction, is superior to other estimators (e.g., OLS with fixed-effects or Standard Difference GMM) in dealing with the highly persistent characteristic of the explanatory variables (Antoniou et al., 2008; Hoechle et al., 2012).

However, using the lagged values of these slow-changing variables as instruments in the System GMM estimation procedure may be questionable. More specifically, if slow-changing explanatory variables are endogenous, then the lagged values of these variables used as instruments will suffer as much from the endogeneity issue as do the current ones (Brown, Beekes, & Verhoeven, 2011). This raises doubts about the validity of the instrumental variables employed in our model. For this reason, we follow the suggestion of Roodman (2009a,b) and carefully test not only the joint validity of the instruments but also the validity of the subsets of System GMM-type instruments as well as standard instruments for the levels equation.

5.2. Preliminary evidence

Table 5 reports the Pearson's correlation matrix for key dependent and independent variables. The correlation coefficient of 0.15 shows that the Tobin's Q is positively related to the percentage of female directors. Although this is only a weak positive linear relationship, it tentatively supports the main hypothesis of our study. The significantly positive relationship between Tobin's Q and the 1-year lagged Tobin's Q is described by the correlation coefficient of 0.58. This supports the proposition that the proper empirical model for the relationship between CG and FP should be considered in a dynamic framework rather than a static one (Wintoki et al., 2012).

Table 4
Between and within standard deviation of corporate governance variables.

Variables	Standard deviations		
	Overall	Between	Within
female	13.76	12.85	5.34
nonexe	20.76	18.93	8.46
dual	0.47	0.42	0.21
lnbsize	0.20	0.18	0.08
block	20.86	18.71	9.11

Note: The notation is as defined in Table 1.

Table 5
Correlation matrix for variables.

	lnq	*female*	*nonexe*	*dual*	*lnbsize*	*block*	*lnfage*	*fsize*	*lev*	*laglnq*
lnq	1.00									
female	0.15***	1.00								
nonexe	−0.07	−0.05	1.00							
dual	0.11**	0.12***	−0.30***	1.00						
lnbsize	0.13***	0.09*	−0.13***	0.10**	1.00					
block	0.13***	−0.07	0.10**	−0.14***	−0.26***	1.00				
lnfage	−0.19***	−0.04	0.14***	−0.04	0.05	−0.04	1.00			
fsize	0.21***	0.04	−0.08*	0.01	0.23***	0.11**	−0.00	1.00		
lev	0.07	−0.09**	−0.11**	−0.10**	0.04	0.10**	−0.14***	0.36***	1.00	
laglnq	0.58***	0.13**	−0.03	0.07	0.12**	0.12**	0.08	0.25***	−0.04	1.00

Note: Asterisks indicate significance at 10% (*), 5% (**), and 1% (***). The notation is as defined in Table 1.

With the exception of the variables *nonexe* and *lev*, the other explanatory variables are significantly correlated with the regressant. From Table 5, the highest significant correlation coefficient among independent variables is 0.36. As suggested by Damodar (2004), unless correlation coefficients among regressors exceed 0.80, multi-collinearity will not be a serious problem for multivariate analysis. Thus, there may be no problem with multi-collinearity among the regressors included in our regression models. In an additional (untabulated)analysis, we conduct a multi-collinearity diagnostic for variables in the model by using variance inflation factors (VIFs). The results show that the highest VIF is 1.73 and the average of VIFs is 1.30, suggesting that multi-collinearity may not be the problem in this study.

The differences in the mean values of Tobin's Q between firms with and without women on their boards are presented in Fig. 1. Intuitively, companies with female directors very likely performed better than those without women on their boards for all years from 2008 to 2011. The clearest evidence was from 2010 when, on average, the Tobin's Q of companies with female directors was 20.40% higher than that of their counterparts. Overall, the graph demonstrates that gender diversity in the boardroom might have a positive relation to firm financial performance, which is consistent with the correlation coefficient between the two variables as reported in Table 5.

Across the full sample, the average value of the Tobin's Q ratio of companies with female board directors was 15.30% higher than those without and the difference was statistically significant at the 1% level (see Table 6 for the *t*-test results). The results reported in Table 6 indicate that the null hypothesis of equal population means[11] should be rejected in the years 2009, 2010, 2011, and across the full sample. It is plausible that in the years when they have female directors on their boards, companies tend to achieve better financial performance measured by Tobin's Q. This finding tentatively supports our main hypothesis that board gender diversity will have a positive impact on firm financial performance. Since the *t*-test procedure does not account for other factors that may interact with the gender diversity–FP relationship, it is difficult to draw causal inferences. The next section presents a further exploration of this relationship through multivariate regression analysis.

5.3. Multivariate regression analysis

Initial multivariate regression results, conducted by using the OLS approach for pooled data, are reported in column 2 of Table 7. It is evident that the percentage of female directors in boardrooms (*female*) is positively related statistically to Tobin's Q at the 1% level ($p = 0.00$), thus providing strong support for our main hypothesis. The coefficient on *female* ($\beta = 0.005$) means that if the percentage of female directors in boardrooms increases by one percentage point, the predicted Tobin's Q will increase, on average, by approximately 0.50%, holding all other factors fixed.

It should be noted that such a percentage change is economically large, given that the size of boards in Vietnamese listed companies ranges between three and eleven members. For example, a change from a board with one woman and seven men to a board with three women and five men leads to a 25 percentage point change. Consequently, the predicted Tobin's Q will increase by approximately $25 \times 0.50\% = 12.50\%$ or, more exactly, by $100 \times [\exp(0.005 \times 25) - 1] \approx 13.31\%$. This finding is consistent with that reported by prior studies including Reddy et al. (2008) in the New Zealand market, Carter et al. (2003) and Adams and Ferreira (2009, p. 305, column 1 of Table 9) in the US market, but contrasts with the findings of Rose (2007) in the Danish market. Such mixed results suggest that further exploration is necessary because the OLS estimator cannot control for potential omitted-variable bias caused by the effects of unobserved features of firms which are invariant over time and/or across firms.

As shown in column 3 of Table 7, when we add firm fixed-effects[12] to address the concern of unobserved heterogeneity, the positive relation between *female* and Tobin's Q is no longer significant ($\beta = 0.002$, $p = 0.47$), which is consistent with the result

[11] The *t*-test procedure is conducted initially to investigate whether there is a significant difference in the performance between companies with and without female directors. To capture both cross-sectional and time variances, we follow Adams and Ferreira (2009) in comparing the means of Tobin's Q not only within the cross section but also across firm-year observations. In order to check the robustness of the results, we follow prior studies (e.g., Chen et al., 2010) and conduct the Wilcoxon rank-sum test for differences in medians of Tobin's Q. The (unreported) results obtained from this non-parametric test show that the conclusions obtained from the *t*-test procedure are robust, even after taking the non-normality of the data into consideration.

[12] The Hausman test for a comparison between the fixed-effects and random-effects models was performed. The null hypothesis that the preferred model is random-effects is rejected, ($\chi^2(9) = 624.10$, $p = 0.00$), suggesting that the fixed-effects estimator should be employed.

Table 6
t-Test for equal population means with unequal variances.

Year	Observations	The average values of Tobin's Q		
		Without female directors	With female directors	Difference
2008	119	0.77	0.84	−0.07
2009	118	0.97	1.15	−0.19**
2010	119	0.80	0.97	−0.16***
2011	116	0.60	0.69	−0.09*
Full sample	472	0.79	0.91	−0.12***

Note: Asterisks indicate significance at 10% (*), 5% (**), and 1% (***).

obtained by Carter et al. (2010) in the US market. This indicates that the significantly positive relation uncovered by the static OLS model may be driven by omitted variable biases. However, it is noteworthy that the fixed-effects approach is implemented under the assumption of strict exogeneity which implies that the CG and control variables are not correlated with the error term in the model. This assumption is criticised for its unreliability when the other sources of endogeneity, namely simultaneity and dynamic endogeneity, very likely arise in the board structure–performance relation in general (Wintoki et al., 2012) and in the gender diversity–performance relation in particular (Adams & Ferreira, 2009). This suggests that while the fixed-effects method in our study produces better estimations, it cannot take into account other potential sources of endogeneity. For this reason, we cannot make reliable causal inferences from the results of model (ii), thus suggesting that the fixed-effects model (within-groups estimators) appears to be undesirable, at least in our study.

In another attempt to capture unobserved heterogeneity, we include a lagged performance variable as an explanatory variable; that is, we move from the static OLS estimator to the dynamic one. As shown in column 4 of Table 7, the estimated coefficient on *female* is statistically different from zero ($\beta = 0.003$, $p = 0.01$), thus providing support for our research hypothesis. Notably, we report that past performance can significantly explain the variation in current performance ($\beta = 0.698$, $p = 0.00$). This is consistent with Wintoki et al. (2012), who showed the importance of using lagged performance variables to control for the dynamic nature of the CG–FP association.

However, the coefficients of CG variables in the dynamic OLS model are considerably smaller than those of the static OLS model. According to Wintoki et al. (2012), such a substantial reduction in the magnitude of the estimated parameters of key CG variables in the dynamic model suggests that the current CG variables are correlated with past FP. This again confirms the dynamic relation between CG and FP. It should be noted that the R^2 in the dynamic model (0.69) is over twice as great as the R^2 in the static model (0.30), indicating that the past performance variable considerably improves the model's power to explain the variation in current performance. Although the dynamic OLS estimator is an improvement over the static models and although our findings are consistent with that of previous studies, they appear to be driven by endogenous biases, such as simultaneity, which cannot be controlled by the pooled OLS method. In addition, the presence of the lagged dependent variable among the regressors makes the OLS estimated parameters biased and inconsistent (Wooldridge, 2002).

Taking into account the concern of the dynamic nature of the board structure–FP relationship, we follow Wintoki et al. (2012) in employing the two-step System GMM estimator. The results reported in the last column of Table 7 show that the percentage of female directors is positively and statistically significantly related to Tobin's Q at the 5% level ($p = 0.04$).[13] The coefficient on *female* ($\beta = 0.021$) means that a ten percentage point increase in the ratio of female directors will, on average, increase the predicted Tobin's Q by approximately 21%, holding all other factors fixed. As mentioned above, this is a strong effect given that the board size of listed companies in Vietnam is small. This result of the System GMM model is consistent with those obtained by using the pooled OLS model and the dynamic OLS model, thereby suggesting that our findings are robust to alternative econometric approaches. This result is also consistent with the findings of several prior studies that confirm the positive relationship between gender diversity and FP (e.g. Campbell & Mínguez-Vera, 2008; Dezsö & Ross, 2012). Our results imply that board gender seems to affect firm value, a point which is in general agreement with Adams et al. (2011, p. 31), who suggest that “shareholders may value female directors because they are better monitors and because they may alleviate value-decreasing stakeholder conflicts”.

In the System GMM model, the coefficient on 1-year lagged Tobin's Q is statistically positive at the 1% level ($\beta = 0.633$, $p = 0.00$), thus suggesting that past performance can help control for unobserved historical factors in the relationship between CG and FP. This empirical evidence strongly supports the arguments of Wintoki et al. (2012), among others, that the link between CG and FP should be examined in a dynamic framework.

Regarding the variable *nonexe*, the results obtained from the pooled OLS, fixed-effects, and dynamic OLS models show that the presence of non-executive directors has no significant impact on FP. However, when moving to the System GMM model, we find that this relationship is significantly negative at the 5% level ($\beta = -.019$, $p = 0.02$). This conclusion is in line with Nowland (2008), which challenged the agency theory's viewpoint regarding the vital role of non-executive directors in monitoring managerial

[13] Based on small-sample corrections, we have reported *t*-test instead of *z*-test statistics for the estimated coefficients, and *F* test statistics instead of *Wald Chi-squared* test statistics for overall fit of the System GMM models.

Table 7
The effect of gender diversity on firm performance.

Regressant: *lnq*	Static models		Dynamic models	
	Pooled OLS	Fixed-effects	Pooled OLS	System GMM
(1)	(2)	(3)	(4)	(5)
constant	−3.038***	−0.272	−0.685*	−4.795*
	[0.00]	[0.88]	[0.06]	[0.10]
laglnq			0.698***	0.633***
			[0.00]	[0.00]
female	0.005***	0.002	0.003**	0.021**
	[0.00]	[0.47]	[0.01]	[0.04]
nonexe	−0.001	−0.001	−0.001	−0.019**
	[0.56]	[0.62]	[0.29]	[0.02]
dual	0.087**	0.119**	0.037	−0.017
	[0.04]	[0.02]	[0.21]	[0.93]
lnbsize	0.204**	0.263*	0.135**	−1.429
	[0.03]	[0.10]	[0.03]	[0.17]
block	0.003***	0.003**	0.001*	0.014**
	[0.00]	[0.01]	[0.09]	[0.03]
lnfage	−0.014	−0.011	0.051	0.430**
	[0.72]	[0.89]	[0.11]	[0.01]
fsize	0.073***	−0.031	−0.005	0.227*
	[0.00]	[0.66]	[0.71]	[0.08]
lev	0.000	0.007***	0.001**	0.000
	[0.97]	[0.00]	[0.05]	[0.99]
Industry dummies	No	No	No	No
Firm fixed-effects	No	Yes	No	Yes
Year dummies	Yes	Yes	Yes	Yes
Number of observations	448	448	352	352
R-squared	0.30	0.59	0.69	
F statistic	15.65	36.56	54.30	12.72
Number of instruments				21
Number of clusters				120
Hansen-J test of over-identification (*p*-value)				0.22

Note: This table presents the results from estimating the Eq. (2) through the use of alternative models, including: (i) a pooled OLS model, (ii) a fixed-effects model, (iii) a dynamic pooled OLS model, and (iv) a dynamic fixed-effects model (System GMM). For models (i) and (ii), it is assumed that $\varphi = 0$. Asterisks indicate significance at 10% (*), 5% (**), and 1% (***). The notation is as defined in Table 1. The *p*-values are presented in brackets. The *p*-values of models (i), (ii), and (iii) are based on robust standard errors corrected for potential heteroskedasticity and serial correlation in the error term. The *p*-values of model (iv) are based on Windmeijer-corrected standard errors. Lags 2 and 3 of the levels of FP variable; lag 2 of the levels of CG and control variables are employed as GMM-type instruments for the differenced equation. Lag 1 of the first differences of FP, CG, and control variables are used as GMM-type instruments for the levels equation. Year dummies and *lnfage* are treated as exogenous variables. Year dummies are unreported.

behaviours and in improving FP. Regarding the other CG variables, we find that there is statistical evidence of a significantly positive link between concentrated ownership and FP ($\beta = 0.014$, $p = 0.03$). This result is consistent in all four models applied in our study and similar to that obtained by Victoria (2006), among others. The positive relationship between block-holder ownership and performance is in agreement with the agency theory perspective that ownership concentration helps to reduce agency problems arising from the separation of ownership and control (Shleifer & Vishny, 1986). This, in turn, is expected to improve performance.

It is obvious from column 5 of Table 7 that the significantly positive relationship between board size and FP, revealed by the static OLS, fixed-effects, and dynamic OLS models, disappears when we control for dynamic endogeneity and simultaneity by using the System GMM model ($\beta = -1.429$, $p = 0.17$). This result accords with the findings of Pham, Suchard, and Zein (2011), Schultz et al. (2010) and Wintoki et al. (2012), who argued that such significant links, estimated by the pooled OLS and fixed-effects models, may be the result of spurious correlations. Similarly, the relation between CEO duality and FP changes from significantly positive to insignificantly negative when we move from the static OLS and fixed-effects models to the System GMM model. This result, once again, supports the argument of Schultz et al. (2010) and Wintoki et al. (2012), among others, that taking the dynamic nature of the relationship between CG and FP into consideration is essential to ensure the reliability of causal inferences.

With regard to the control variable *leverage*, we find that the positive relationship between financial leverage and FP uncovered by the fixed-effects model and the dynamic pooled OLS model disappears when the potential sources of endogeneity are taken into consideration (column 5 of Table 7). Several robust-checking models reported in Table 9 and Table 10 also confirm that the estimated coefficient on *leverage* is not statistically different from zero at any conventional levels of significance, suggesting that financial leverage has no impact on FP. Although this finding is consistent with that of Nguyen et al. (2014) and Schultz et al. (2010), among others, the relationship between financial leverage and FP is not really clear in practice. The discussion below provides some possible explanations for our finding.

A recent study undertaken by Jiraporn et al. (2012) suggested that debt financing and CG mechanisms may substitute for each other to alleviate agency cost whereby FP is improved. If that is the case, it is plausible to argue that the potential performance effect of financial leverage in Vietnamese companies is likely to be replaced by the stronger effects of other CG mechanisms, including

Table 8
Difference-in-Hansen tests of exogeneity of instrument subsets.

Tested instrument subsets	Test statistics (Chi-squared)	Degrees of freedom	*p*-Value
Panel A: System GMM-type instruments			
Instruments for the levels equation as a group	10.82	8	0.21
lnq_{it-2} and lnq_{it-3} (for the transformed equation)	1.36	2	0.51
Δlnq_{it-1} (for the levels equation)	1.55	1	0.21
Instruments for board structure variables	11.60	8	0.17
Instruments for control variables	10.56	6	0.10
Panel B: Standard instruments			
2009 and 2010 *year dummies*, and *lnfage*	6.25	3	0.10

Note: This table presents difference-in-Hansen tests of the exogeneity of instrument subsets, under the null hypothesis of joint validity of a specific instrument subset. The test statistics are asymptotically chi-squared distribution with degrees of freedom equal to the number of questionable instrumental variables (Roodman, 2009a). GMM instrument subset used for the levels equation includes one-year lagged differences of FP, board structure, and control variables (Δlnq_{it-1}; $\Delta female_{it-1}$; $\Delta nonexe_{it-1}$; $\Delta dual_{it-1}$; $\Delta lnbsize_{it-1}$; $\Delta block_{it-1}$; $\Delta fsize_{it-1}$; and Δlev_{it-1}). GMM instrument subset used for board structure variables includes lag 1 of the first difference and lag 2 in levels of board structure variables ($female_{it-2}$ and $\Delta female_{it-1}$; $nonexe_{it-2}$ and $\Delta nonexe_{it-1}$; $dual_{it-2}$ and $\Delta dual_{it-1}$; $lnbsize_{it-2}$ and $\Delta lnbsize_{it-1}$). GMM instrument subset used for control variables includes lag 1 of the first differences and lag 2 in levels of control variables ($block_{it-2}$ and $\Delta block_{it-1}$; $fsize_{it-2}$ and $\Delta fsize_{it-1}$; lev_{it-2} and Δlev_{it-1}). 2008 and 2011 year dummies are dropped due to collinearity.

ownership concentration (measured by *block*) and board diversity (measured by *female*). In consequence, the estimated coefficient on *leverage* should not be statistically different from zero.

In a similar vein, González (2013) argued that the association between financial leverage and FP is likely to be contingent upon two contradictory antecedents: (i) the cost of financial distress; and (ii) the benefits of the disciplinary role of debt financing. A firm with higher financial leverage may suffer from higher costs of financial distress but may also benefit from the disciplinary role of debt financing, by which managers are forced to take value-maximising decisions (González, 2013). Therefore, the net effect of financial leverage on FP can be neutralised if neither of these two antecedents is predominant.

As reported in the last row of Table 7, the Hansen-J test[14] yields the *p*-value of 0.22, suggesting that the null hypothesis that our instruments are valid cannot be rejected at any conventional levels of significance. We also follow the recommendation of Roodman (2009b) about good practices in implementing System GMM estimation and apply the difference-in-Hansen test to the subsets of System GMM-type instruments, as well as standard instrumental variables for the levels equation. Table 8 presents difference-in-Hansen tests of the exogeneity of instrument subsets, under the null hypothesis of joint validity of a given instrument subset. We find no statistical evidence to reject the null hypothesis, suggesting that the subsets of instruments are econometrically exogenous.

5.4. Robustness checks

5.4.1. The impact of macroeconomic environment on firm performance[15]

Macroeconomic variables, such as GDP growth, may have potential impacts on firm value. Koller et al. (2010) posited that the value of firms may reflect the expectations of investors about overall macroeconomic activities since firm valuation is directly influenced by the assumptions of major macroeconomic variables. More specifically, the changes in Tobin's Q ratio over time may partly reflect the variations in the valuation of future growth opportunities which arise from exogenous economic conditions (Pham et al., 2011). Following Jiang et al. (2012), we use GDP growth measured by the annual percentage growth rate of GDP calculated at market prices as a proxy for the macroeconomic environment in which firms operate. In line with the result of Jiang, Feng, and Zhang (2012), we find that GDP growth is positively associated with Tobin's Q at the 1% level of significance ($\beta = 0.45$, $p = 0.00$). The results reported in column 2 of Table 9 show that the estimated coefficients on the variable of interest (*female*) and on the other CG variables are generally unchanged. Therefore, our findings remain robust even after controlling for the potential impact of macroeconomic environment on FP.

5.4.2. The impact of alternative proxies for gender diversity on firm performance

In order to check the robustness of the estimations reported in column 5 of Table 7, we follow Campbell and Mínguez-Vera (2008) and employ two alternative proxies for gender diversity, including a gender diversity dummy variable (*d1women*) and the Blau index

[14] Because the validity of System GMM estimation is strongly influenced by the validity of instrumental variables, it is very important to diagnose whether or not the instruments are exogenous. According to Arellano and Bond (1991), there are three alternative tests for the validity of instrumental variables, namely: (i) the Arellano–Bond test for second-order serial correlation [the AB AR(2)] in the first differences of residual series; (ii) the Hansen-J test of over-identifying restrictions; and (iii) the Hausman specification test. Among them, we use the Hansen-J test in this study for the two following reasons. First, since the AB AR(2) test statistic is only defined if min $T \geq 5$ (Arellano & Bond, 1991), this test cannot be used in our circumstance in which $T = 4$ years. Second, while the power of the Hausman specification test is questionable, especially in the presence of outliers (Arellano & Bond, 1991), the Hansen-J test is considered a standard test for the joint validity of the instrumental variables after GMM estimation (Baum, 2006; Roodman, 2009a).

[15] It is worth noting that we employ year dummy variables in our base-line model to account for time-specific effects (e.g., inflation rate, demand shocks, and other macroeconomic conditions) which are common to all companies and can change through time. The impact of macroeconomic variables on performance (if any), therefore, is already included in these year dummies. This subsection aims to examine in more detail the potential effect of GDP growth, one of the most important macroeconomic variables, on FP. This helps to check the robustness of our key findings when the potential impact of the macroeconomic environment on FP is explicitly taken into consideration.

Table 9
Robustness checks with *gdpgrow* and alternative proxies for gender diversity.

Regressant: *lnq*	System GMM		
	With *gdpgrow*	With *d1women*	With *blau*
(1)	(2)	(3)	(4)
constant	−7.560**	−4.867	−4.560
	[0.02]	[0.13]	[0.10]
laglnq	0.610***	0.629***	0.607***
	[0.00]	[0.00]	[0.00]
female	0.022**		
	[0.03]		
d1women		0.379*	
		[0.08]	
blau			1.461**
			[0.02]
nonexe	−0.020***	−0.016***	−0.018***
	[0.01]	[0.01]	[0.01]
dual	0.004	−0.117	−0.078
	[0.98]	[0.48]	[0.66]
lnbsize	−1.345	−1.051	−1.368
	[0.14]	[0.22]	[0.15]
block	0.013**	0.009**	0.012**
	[0.02]	[0.05]	[0.02]
lnfage	0.435***	0.275**	0.378**
	[0.01]	[0.03]	[0.01]
fsize	0.228*	0.228	0.222*
	[0.08]	[0.15]	[0.10]
lev	0.000	−0.007	−0.003
	[0.96]	[0.43]	[0.74]
gdpgrow	0.450***		
	[0.00]		
Number of observations	352	352	352
F statistic	14.27	14.56	14.23
Number of instruments	20	27	20
Number of clusters	120	120	120
Hansen-J test of over-identification (*p*-value)	0.23	0.30	0.23

Note: This table presents the robust results from estimating modified Eq. (2) through the use of an additional control variable (*gdpgrow*), and two alternative proxies for gender diversity, including: (i) *d1women*, and (ii) *blau*. The definitions of the variables are provided in Table 1. Asterisks indicate significance at 10% (*), 5% (**), and 1% (***). The *p*-values are based on Windmeijer-corrected standard errors and presented in brackets. Lag 2 of the levels of FP, CG and control variables are employed as GMM-type instruments for the differenced equations of the models in columns (2) and (4). Lag 2 of the levels of FP variable, lags 2 and 3 of CG and control variables are employed as GMM-type instruments for the differenced equations of the model in column (3). Lag 1 of the first differences of FP, CG, and control variables are used as GMM-type instruments for the levels equations in all of the models. Year dummies, *gdpgrow*, and *lnfage* are treated as exogenous variables. Year dummies are not reported.

for gender (*blau*). The finding reported in columns 3 of Table 9 shows that the presence of female directors in the boardroom (measured by *d1women*) is positively related to firm value at the 10% level ($p = 0.08$). The coefficient on *d1women* ($\beta = 0.379$) implies that the difference in the predicted Tobin's Q between companies with at least one female director on their boards and those without is about 37.90% or, more exactly, $100 \times [\exp(0.379) - 1] \approx 46\%$. Similarly, the result from column 4 of Table 9 indicates that heterogeneous boards (measured by *blau*) have a statistically positive impact on FP at the 5% level ($\beta = 1.461$, $p = 0.02$). Thus the positive relationship between board gender diversity and FP remains robust when alternative proxies for gender diversity are employed.

To capture the potential effect of the number of female directors, we include in Eq. (2) one dummy variable that takes a value of one if there are at least two female directors and zero otherwise (denoted as *d2women*). As reported in column 2 of Table 10, we find that the estimated coefficient on *d2women* ($\beta = 0.610$) is statistically significant at the 5% level and considerably larger than that on *d1women* reported in column 3 of Table 9 ($\beta = 0.379$). This finding suggests that boards with at least two female directors appear to have a stronger effect on FP than those with at least one. This empirical result generally supports the perspective of ‘critical mass theory’ proposed by Kanter (1977) that women may have a more significant effect on a group when they increase from a token number to form a significant minority of the group. In other words, if female board representation increases board effectiveness and FP, then that effect should be more pronounced when the number of female directors increases (Liu et al., 2014). However, given the significantly positive coefficient on both *d1women* and *d2women*, we also support the view suggested by Zaichkowsky (2014), that although two or more women on boards appear to have a stronger effect on firm outcomes, even one woman can make a difference.

It is noteworthy that although the relationship between board gender diversity and FP appears to be significantly positive, it is not necessarily a linear relationship. To check empirically for possible non-linearity in the board gender diversity–performance relationship, a quadratic term of the variable *female* (denoted as *female_squared*) is included in Eq. (2). In an un-tabulated analysis, we apply the pooled OLS approach to the modified Eq. (2) and find that: (i) the estimated coefficient on *female_squared* is

Table 10
Robustness checks with alternative proxies for gender diversity (cont.).

Regressant: *lnq*	System GMM	
	With *d2women*	With *female_squared*
(1)	(2)	(3)
constant	−5.261**	−4.120**
	[0.03]	[0.02]
laglnq	0.602***	0.545***
	[0.00]	[0.00]
d2women	0.610**	
	[0.05]	
female		0.033**
		[0.04]
female_squared		−0.001
		[0.14]
nonexe	−0.018**	−0.012*
	[0.04]	[0.06]
dual	0.009	−0.127
	[0.96]	[0.46]
lnbsize	−0.973	−0.588
	[0.30]	[0.37]
block	0.012**	0.008*
	[0.03]	[0.06]
lnfage	0.374**	0.229*
	[0.02]	[0.10]
fsize	0.224*	0.166*
	[0.05]	[0.05]
lev	−0.000	−0.003
	[0.97]	[0.51]
Firm fixed-effects	Yes	Yes
Year dummies	Yes	Yes
Number of observations	352	352
F statistic	15.76	14.84
Number of instruments	22	30
Number of clusters	120	120
Hansen-J test of over-identification (*p*-value)	0.15	0.16

Note: This table presents the robust results from estimating modified Eq. (2) using the two-step System GMM approach. Column (2) presents the robust results when we add one new dummy variable (denoted as *d2women*) to Eq. (2) to capture the potential effect of the number of female directors. Column (3) presents the robust results when we add a quadratic term of *female* (denoted as *female_squared*) to Eq. (2) to empirically check for the possible non-linearity in the board gender diversity–performance relationship. The definitions of the variables are provided in Table 1. Asterisks indicate significance at 10% (*), 5% (**), and 1% (***). The *p*-values are based on Windmeijer-corrected standard errors and presented in brackets. Year dummies are not reported.

statistically insignificant ($\beta = -0.0001$; p-value $= 0.142$); and (ii) the estimated coefficient on *female* is still significantly positive ($\beta = 0.0060$; p-value $= 0.021$).

To further challenge these results, we apply the two-step System GMM estimation approach on the modified Eq. (2) and achieve similar results. Specifically, as reported in column 3 of Table 10, the estimated coefficient on *female_squared* is statistically insignificant ($\beta = -0.001$; p-value $= 0.14$), whereas the coefficient on the variable of interest (*female*) is still significantly positive ($\beta = 0.033$; p-value $= 0.04$). The results obtained from the OLS and System GMM methods allow us to conclude that there is not enough statistical evidence to support a non-linear association between board gender diversity and the performance of Vietnamese companies.

Nevertheless, one concern is that over-diversification will wipe out the variety and/or the balance of board composition, so that gender diversification leading to an all-female BOD may be counterproductive. The empirical question here is: what is the breakpoint at which an undesired effect of gender diversification occurs? To provide a quick look at the relationship between FP and board gender diversity which may help to find a possible answer to this question, we perform a median-band plot[16] together with a scatter-plot for *Tobin's Q* against the *Blau index*. The *Blau index* for gender is employed since, as mentioned earlier in Section 4.2, it allows us to capture both aspects of diversity, including gender variety and gender balance. As shown by the median spline in Appendix 2 the medians of *Tobin's Q* increase with the medians of the *Blau index* until the latter reaches about 0.30 and then seem to remain unchanged when the *Blau index* goes beyond 0.30. This suggests that 0.30 is likely to be the breakpoint at which the undesired effect of gender diversification may occur.

To check this result empirically, we carry out a segmented regression analysis in which the sample is divided into two separate datasets on the basis of the *Blau index*. Accordingly, we rerun Eq. (2) on the sub-dataset with a *Blau index* smaller than 0.30, and on the other with a *Blau index* equal to or larger than 0.30. The results reported in Table 11 show that the relationship between *Tobin's Q* and the *Blau index* appears to change over different intervals of the *Blau index*. More specifically, for firms with a *Blau index* smaller than 0.30, the *Blau index* is significantly positively related to financial performance (columns 2 and 3 of Table 11). By contrast, for firms with a *Blau index* equal to or larger than 0.30, the relationship becomes insignificant (columns 4 and 5 of Table 11). These results

[16] A related two-way median spline providing a smoother version of the median-band plot is included in Appendix 2.

Table 11
Robustness checks using segmented regression analysis.

Regressant: *lnq*	Blau Index < 0.3		Blau Index ≥ 0.3	
	OLS	System GMM	OLS	System GMM
(1)	(2)	(3)	(4)	(5)
…	…	…	…	…
blau	0.550***	1.493**	0.215	0.265
	[0.00]	[0.04]	[0.58]	[0.53]
…	…	…	…	…
Number of observations	209	209	143	143
R-squared	0.70		0.70	
F statistic	31.90	10.80	26.46	14.10
Hansen-J test (*p*-value)		0.43		0.47

Note: This table presents the estimated coefficient on *blau* obtained from segmented regression analysis in which the sample is divided into two separate datasets on the basis of the *Blau index*. Accordingly, Eq. (2) is estimated on the sub-dataset in which the *Blau index* is smaller than 0.30, and on the other sub-dataset in which the *Blau index* is equal to or larger than 0.30. Columns (2) and (4) present the results obtained from the pooled OLS approach. Columns (3) and (5) present the results obtained from the two-step System GMM approach. The definitions of the variables are provided in Table 1. Asterisks indicate significance at 5% (**), and 1% (***). The *p*-values are presented in brackets. The *p*-values reported in columns (2) and (4) are based on robust standard errors corrected for potential heteroskedasticity and serial correlation in the error term. The *p*-values reported in columns (3) and (5) are based on Windmeijer-corrected standard errors. To save space, the estimated coefficients on other variables are not reported.

remain robust when alternative econometric techniques are applied and consistent with what we can observe from the median-band plot, that there is likely to be an upward trend in Tobin's Q as the *Blau index* increases to 0.30. After this point, there is no further significant trend in the Tobin's Q.

The critical *Blau index* of 0.30 can be approximately translated into two critical percentages of female directors: ether 20% or 80%. However, it is impractical to consider the critical percentage of 80%, given that the maximum proportion of female directors on boards in our sample is just about 67%. Consequently, it is evident from our empirical analysis that a *Blau index* of about 0.30, corresponding to a ratio of about 20% of women on the BOD, is the breakpoint at which the potential performance effect of female board representation may change.[17] Although we are unable to answer explicitly, for the purposes of the current study, what the mechanism behind the scene is, we suggest one possible explanation for our finding that greater gender diversity on boards will add value as long as the benefits obtained from the diversification are not outweighed by its costs. We believe that the trade-off between the costs and benefits of board gender diversification may offer insight into developing a theoretical framework that can provide a clear-cut prediction about the nature of the board gender diversity–FP association.

6. Conclusion and limitations

Over the last decade, the relationship between board gender diversity and FP has received considerable attention from scholars and the mass media worldwide. However, this topic has not been explored in Vietnam. To the best of our knowledge, this research makes the first effort to discover the relationship between gender diversity and FP of the publicly listed companies in this country. After controlling for firm size, firm age, time (year), leverage, ownership structure, unobserved historical factors, and other CG characteristics, we report that gender diversity in the boardrooms of the publicly listed companies in Vietnam tends to have a positive effect on financial performance measured by Tobin's Q.

In addition, we find that the number of female directors in the boardroom also makes a difference. Boards with at least one female director seem to outperform those with none and boards with at least two female directors appear to have a stronger effect on FP than those with at least one. We further document that the nature of the relationship between board gender diversity and FP may change when the percentage of women reaches the breakpoint of 20%. Theoretically, our findings are in agreement with the perspectives of agency theory and resources dependence theory regarding the positive influences of board diversity on board effectiveness, which in turn will result in improved FP. Empirically, our findings hold even after taking into account dynamic endogeneity, simultaneity, and unobserved time-invariant heterogeneity inherent in the CG–FP relationship.

The contribution of this paper to the extant literature is twofold. First, by providing robust evidence from a transitional economy characterised by poor CG practices, our study supports the view proposed by Adams and Ferreira (2009) and Gul et al. (2011) that boards with greater gender-diversity provide additional monitoring and act as a substitute governance mechanism from which weak-governed companies may benefit. Although our findings are in no way intended to support mandating a gender quota system for the BOD, they do offer an important implication for policy formulation. We suggest that efforts to rejuvenate CG by increasing the number of women on Vietnamese boards of directors (and perhaps BODs in other Asian markets sharing similar CG characteristics) should take the existing conditions of the CG system into consideration.

Based on the research context, furthermore, our findings may also imply that in Asian developing countries where women are traditionally subordinate to men, female directors have the potential to add value if they enjoy a supportive institutional environment and if

[17] In order to check for robustness, we repeat the segmented regression procedure in which the sample is divided into two datasets on the basis of *female*. Accordingly, we rerun Eq. (2) on the sub-dataset with *female* less than 20%, and on the other with *female* equal to or greater than 20%. We find that the results (unreported due to space limitations) are not qualitatively different from those reported in Table 11.

the advancement of women is consistently promoted. Therefore, we suggest that a better institutional environment for women plays an important role in board gender diversification, which may have a positive effect on FP. Second, this study contributes to the international debate on the gender diversity–performance relationship by adopting a well-structured dynamic modelling approach to address the potential endogeneity concerns inherent in the CG–FP association, thus helping to confirm and strengthen prior research findings.

Despite the above-mentioned contributions, this study does have some limitations, many of which may indicate fruitful avenues for future research. First, given that the variables relating to board structure change slowly over time, potentially vitiating panel data estimations, this study's 4-year dataset may render comprehensive explanations of governance–performance dynamics ineffective. Following Wintoki et al. (2012), we suggest that datasets covering a longer period of time may enable future research to overcome the highly persistent feature of board structure variables by using data at 2-year intervals rather than annually.

Secondly, this research considers only the relationship between gender diversity on the BOD and FP. Given the two-tier board structure of listed companies in Vietnam, it might be useful for future research to treat gender diversity on the BOS as a factor driving the firms' profitability. Moreover, like most previous empirical studies on CG, our sample selection process relies primarily on the availability of data, including firm annual reports and corresponding financial reports. It is likely that the selected firms are the more transparent ones and, therefore, could actually be well governed and/or better performing firms. If that is the case, our research will suffer from selection bias which hinders interpretation and generalisation.

Finally, although this study does find a significant link between board gender diversity and the performance of Vietnamese listed companies, the channels through which female directors positively affect financial performance remain unclear. It is argued that if the presence of women on the BOD matters for firm outcomes, then there may be gender-based differences in behaviour and characteristics between female and male directors (Mohan, 2014). We believe that understanding the personality traits of female directors, such as consensus-building ability, management style, or attitude to risk, is essential to shed light on the potential channels through which director gender matters for FP.

Additionally, there is empirical evidence that greater gender diversity on boards may promote better opportunities for women to be appointed to top management positions. Female representation on top management teams, in turn provides a feminine managerial expertise and helps to improve managerial task performance (Dezsö & Ross, 2012). If this evidence is considered, female directors may also add value through their contribution in choosing the CEO and motivating women in senior management positions. For this reason, we believe that investigating the role of women at top management levels in interaction with the role of women on boards offers potential for understanding another channel through which female directors add value.

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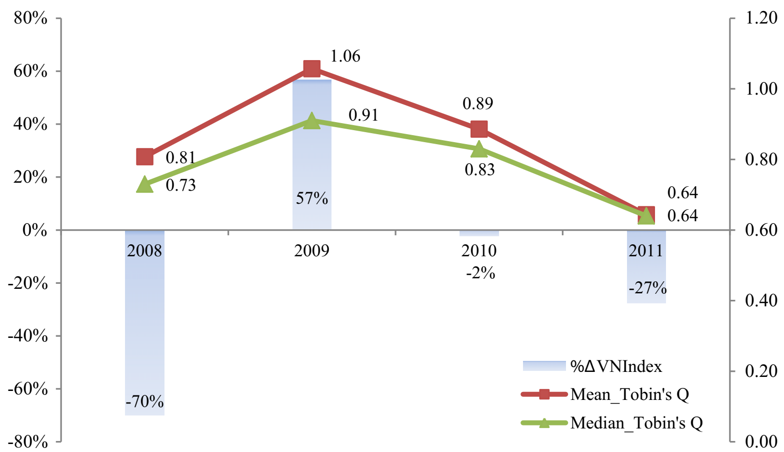
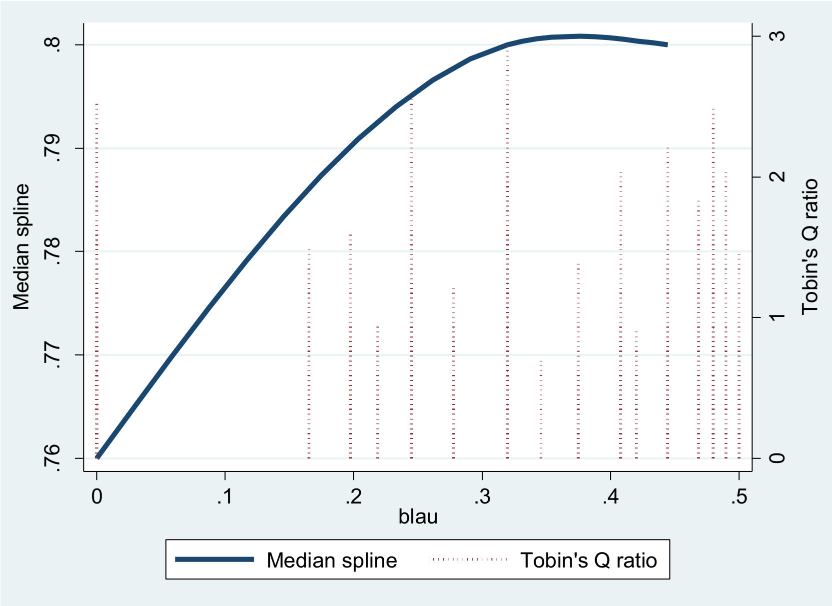
Appendix A. Supplementary data

Supplementary data to this article can be found online at http://dx.doi.org/10.1016/j.iref.2014.11.022.



Appendix 1. Changes in the Vietnamese stock index (%) vs. Tobin's Q, 2008–2011 Source: %ΔVN-index and the mean and median values of Tobin's Q are based on data collected from the HOSE website and Thomson One Banker (Worldscope database).



Appendix 2. The median-spline plot and scatter-plot for *Tobin's Q* against the *Blau index*.

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