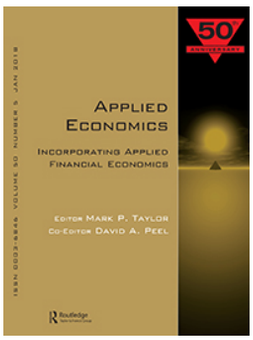

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A synthetic cohort analysis of female labour supply: the case of Thailand

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ABSTRACT

This study analyzes trends and patterns in employment, hours worked, and women's wages for Thai women, and relations among these variables in models of female labour supply. Labour supply behaviour of Thai women is investigated with synthetic cohort data defined by age, year of birth, and level of educational attainment, constructed from annual labour force surveys from 1985 to 2004. According to pseudo-panel estimates, wage increases lead to a reduction in hours worked, but also an increase in the employment/population ratio, with elasticity estimates that are robust across a variety of specifications. Estimates based on disaggregation by marital status show that marriage provides protection against wage decreases, allowing married women greater choice between hours of work and other household activities. Alternative disaggregations find that younger cohorts of women and those with the highest education show the greatest (positive) response of the employment rate to wage changes, leading to a future Thai labour force with more women and higher educational attainment.

KEYWORDS

Female labour supply; synthetic cohort; Thailand

JEL CLASSIFICATION

C23; J21; J22

I. Introduction

In recent decades, Thailand has experienced substantial feminization of the workforce. Thai society has adopted modern attitudes towards the role of women that have encouraged high levels of female labour force participation. The Thai female participation rate is around 65% whereas the range of the other countries in Southeast Asia, e.g. Malaysia and the Philippines, is around 25–50% (ILO 2008). The earnings gap between men and women has closed markedly: from 34% in 1985 to 9% in 2005 (Nakavachara 2010). These changes in the Thai labour market reflect, in part, increases in schooling and the elimination of gender gaps in primary and secondary schooling (Knodel 1997). With higher education levels come higher wages (Warunsiri and McNown 2010), and external benefits in the form of healthy childbearing and childrearing (Schultz 1995).

Female labour market activity is an important indicator of the economic and social status of women (Mammen and Paxson 2000). The goal of this study is to analyze trends and patterns in female labour force participation, hours worked, and women's wages in Thailand and the relations among these variables in models of female labour supply from 1985 to 2004. This period spans episodes of major change in the Thai economy, including rapid industrialization and growth during the first ten years, followed by the Asian financial crisis of 1997, and the resumption of moderate economic growth following the crisis.

Rapid industrialization, the promotion of female education, integration of women into the modern workforce, and the closing wage gap between men and women make Thailand an interesting case for the study of female labour supply. Empirical studies of female labour supply generally consider either the extensive margin (the female labour force participation rate) or the intensive margin (the association between wages and working hours). Despite several studies on female labour supply in developed countries and developing countries (see comprehensive review in Killingsworth and Heckman 1986; Psacharopoulos and Tzannatos 1989; Assaad and El-Hamidi 2002), there is only limited literature on female labour supply in Thailand including: Schultz (1990), Mammen and Paxson (2000), and Nakavachara (2010). However, these studies do not consider both intensive and extensive margins, nor do

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they analyze the patterns of female labour supply over time despite the existence of significant changes.

Average wages of female workers have generally trended upwards over the twenty year period of this study (1985–2004), although with a flat period between 1993 and 2000 (Table 2); also weekly hours worked by women have increased steadily, except for the crisis year of 1997 (Table 3). However, the ratio of employed women to the total number of women of working age has shown no distinct trend (Table 4). It is difficult to see any systematic associations among these three aggregates that are suggestive of labour supply relations, and this raises interesting questions that have not been addressed in previous studies.

Three main questions are the focus of this article:

(1) Is the response of hours worked to wages positive or negative? And what is a point estimate of this elasticity?
(2) Does an increase in wages lead to an increase in hours worked when considering both intensive and extensive margins?
(3) How do these elasticities vary across different demographic groups, classified by education level, birth cohort, age, urban-rural areas?

This study reassesses these general tendencies with synthetic panel data constructed from repeated cross-sections from Thailand's Labour Force Survey (LFS) collected over the twenty years from 1985 to 2004. This long survey record is unusual for developing countries and provides the opportunity for the use of the synthetic cohort approach initiated by Deaton (1985), with advantages for female labour supply estimation as suggested by Devereux (2007). Following Pencavel's (1998) application to U.S. data, the cross sectional surveys are aggregated into age-education-cohort group means, reducing the problems of measurement error (Antman and McKenzie 2007) and individual heterogeneity that plague individual cross sectional regressions. Furthermore, this approach can capture varying labour market behaviour over time and across generations that cannot be observed with cross-sectional data alone. The usual econometric problems characteristic of labour supply studies – endogeneity of wages and labour supply, sample selection into the labour force – are addressed through instrumental variables techniques and construction of an additive selection term (Pencavel 1998) using the inverse Mills ratio. This approach was proposed by Indrawati and Leuthold (1992) for female labour supply estimation, and by Mora and Mura (2007) specifically for pseudo-panel data. Acknowledging these advantages of the pseudo-panel approach, the study reported here is the only analysis of Thai female labour markets that uses this methodology.

Elasticities are estimated with various choices of control variables, applying both instrumental variables and classical regression approaches, and the range of estimates for the entire sample of women is quite narrow. The robustness of these estimates across various specifications provides confidence in the methodology employed and values of the labour supply elasticities presented here. Disaggregations by urban-rural residence, educational attainment, age, and birth cohort provide extensions of the basic framework.

The next section introduces the underlying theoretical framework and reviews previous studies of female labour supply in developing countries and Thailand in particular. Section III presents the data set and variables used in the estimation; econometric issues, including the identifying assumptions, are discussed in Section IV. Section V analyzes the results, and Section VI concludes.

II. Theory and empirical studies of female labour supply in developing countries

According to the individual labour supply model, the response of labour supply to wages depends on the labour leisure trade-off. For those already working, a change in wages induces both income and substitutions effects, with opposite consequences for working hours. The basic static labour supply model cannot predict whether the slope of labour supply will be upwards or downwards (Killingsworth 1983). Killingsworth and Heckman (1986) conclude that the sign and the magnitude of estimated wage elasticity of female labour supply across studies depend on the data source, the methodology, as well as the demographic group under study.

While the standard expectation is that substitution effects dominate income effects, so that the elasticity of hours worked with respect to wages should be positive, evidence from several developing countries refutes this position (Psacharopoulos and Tzannatos 1989; Schultz 1990; Sharif 1991; Yamada 2008; Dessing 2002).

On the extensive margin there is no income effect, and female labour force participation rates are expected to respond positively to wage changes (Psacharopoulos and Tzannatos 1989; Schultz 1990). Several studies use microdata to study female labour markets in other developing countries; however, the existing studies in developing countries mostly focus on the female labour force participation rather than the work-wage association (Assaad and El-Hamidi 2002).

In addition to wage rates, several other factors determine the female participation rate in the labour market. For example, Fakih and Ghazalian (2015) investigated which firm-related factors and which national factors determine female employment in the Middle East and North African (MENA) region by using the firm-level data from the World Bank's Enterprise Surveys and employing the fractional logit model in the estimation. Note that the MENA region has special characteristics, with low female employment due to social and cultural norms that impede women from working outside home. Their results show that firm-related factors, such as private foreign ownership, export activities, firm size, and labour composition, affect female employment in manufacturing firms; national factors affecting female employment include economic development and gender equality. Their results are consistent with the study by Ahmed, Feeny, and Posso (2016), which shows that firm size and export orientation activities are primary factors affecting female employment in firms in manufacturing sectors of Bangladesh.

In the study of the work-wage association in developing countries, Yamada (2008) finds negative wage elasticities for women in urban areas of Peru, as does Dessing (2002) for the Philippines. Dessing (2002) summarizes other studies that also find negative wage elasticities, and emphasizes that these responses are particularly strong for secondary workers. The theoretical framework behind the hypotheses in the current study is based on the theory presented in Dessing (2002). In addition to time constraints and budget constraints in the basic labour supply model, subsistence constraints are also included, such that an individual chooses between income and leisure subject to all three constraints in order to maximize utility. In the case of low-income workers, the subsistence constraints will limit their choice between leisure and work. With the time constraints and subsistence constraints, this individual will choose to supply labour to maximize the utility given the subsistence wage. If the wage rate in the market equals the subsistence wage, individuals must work full time in order to meet their subsistence needs.

Dessing (2002) presents formal theoretical argument for the dominant income effect that explains the common finding of an inverse relation between wages and hours worked by women. In her model, primary workers are always fully employed, and secondary workers, primarily married women, adjust their hours of work as necessary to maintain a subsistence level of household income. When wages fall, these women increase their hours of work, adding to the already long hours they spend on household chores and child rearing. When wages rise, secondary workers choose to reduce their working hours so that their total hours of household work and employment become more manageable. This dominant income effect is particularly strong at income levels that are near subsistence, consistent with the findings for other low income countries or regions cited above. Dessing's theory is consistent with additional studies in other developing countries that also find a decline in wages drives women to increase their labour supply e.g. in Bangladesh (Sharif 1991), Mexico (Licona 2000), Egypt (El-Hamidi 2003), and India (Dasgupta and Goldar 2006). One objective of the current study is to re-examine this finding with careful attention to econometric issues of labour supply estimation.

The relations between wages and the two measures of labour supply are expected to vary across demographic and geographic categories. For example, female workers in rural areas may face a weaker labour-leisure trade-off compared with urban workers, since child-rearing need not conflict with agricultural work. Marital status may also affect the elasticity of labour supply, as the husband's income provides a form of financial support for the household should a married woman decide to reduce her hours of work. In addition, across different levels of educational attainment and different birth cohorts representing varying levels of modernization, one also expects different labour supply relations working through changing preferences towards labour, leisure, and child rearing. In this study, female labour supply elasticities are estimated under

alternative disaggregations designed to uncover these hypothesized effects.

Several prior studies employ survey data to investigate female labour markets in Thailand. Mammen and Paxson (2000), working with survey data from India and Thailand, investigate the determinants of female labour force participation and how this has changed over time. Lacking wage data, they do not estimate labour supply functions, but focus on the relation between schooling and labour force participation and differences in this relation across birth cohorts and demographic groups. They find female LFP to be lower in cities compared with rural areas. Also, participation rates are lower for recent birth cohorts, a finding they explain as a response to higher household incomes from their husbands. Surprisingly, women with secondary schooling have lower participation rates than those with lower and higher levels of educational attainment.

Schultz (1990) estimates models of participation and hours worked for males and females with the goal of testing the restrictions of the unitary household decision model against a more general bargaining model of the allocation of unearned income. Using Thai household survey data from 1981, he soundly rejects the restrictions of the neoclassical model. He also finds negative wage responses for hours worked by both men and women, employing Heckman selection and instrumental variables estimation strategies.

Although Schultz (1990) also estimates both participation and working hours equations, the study reported here offers several further contributions. Schultz's estimates are based on a single year; here the estimation uses data spanning twenty years of Thai industrialization. Estimates based on samples disaggregated by demographic groups reported in this article show how labour supply elasticities vary as the composition of the population changes with economic and social development. These estimates are based on an alternative methodology that minimizes several econometric issues common to micro-data analysis.

Nakavachara (2010) documents the dramatic changes in education, employment and earnings experienced by Thai women over the past twenty years. Women now outpace men in their educational attainment. Earnings for both males and females grew steadily between 1985 and 1996, dipped during the Asian financial crisis of 1997, and have since recovered and levelled off until the end of her data in 2005. More significantly, the earnings gap between the sexes has shrunk from 34% in 1985 to 9% since 2001. Reducing the gender wage gap reflects the country's economic development (see Cuberes and Teignier 2014 for comprehensive theoretical and empirical review of the linkage between gender inequality and economic growth). According to Nakavachara's decomposition of the earnings gap, observable factors[1] would predict that women's wages should exceed men's wages, primarily due to the higher educational attainment of Thai women today. Although her study does not directly address the issues of labour supply, t highlights the trends in the status of women that play an important role in the market for female workers in Thailand.

To add to this literature, the current study considers both intensive and extensive margins of female labour supply in Thailand, estimating labour supply elasticities for various demographic groups and cohorts of Thai women. Recognizing the potential biases in estimation with individual data, the analysis departs from these prior studies by employing pseudo-panels of synthetic cohorts to deal with well-established problems of individual heterogeneity, wage endogeneity over the life cycle, and sample selection.

III. Data and variables

The data were collected by the National Statistical Office of Thailand (NSO), Statistical Forecasting Bureau, as part of the National Labour Force Surveys (LFS) for 1985–2004. A change in the survey instrument in 2005 precludes the use of data after 2004 in construction of pseudo-panels employed in this study. This study uses the 3rd quarter of each survey year in order to control for the effect of seasonal agricultural labour movements. Thai agricultural workers migrate to work in the cities during the dry season, but return home for the third quarter rainy season (Sussangkarn and Chalamwong 1996).

A woman is classified as employed in this study if she (1) worked during the survey week, (2) reported

[1]In Nakavachara's study, the observable determinants of wages are education, age, marital status, and community and region characteristics.

a number of working hours during the previous week, and (3) reported monthly earnings. The sample includes all employment statuses including private employees, government employees, state-enterprise employees, and the self-employed.

Following Pencavel (1998), this study adopts a pseudo-panel approach (Deaton 1985), constructing group means of all variables from 109,095 individual observations on women sampled from the 20 years of the survey. The youngest woman in the study is 25 and the oldest is 60, avoiding complications caused by schooling or retirement decisions. The groups are defined by age, year of birth, and level of educational attainment. Seven five-year age categories and eight five-year or ten-year birth-year cohorts (starting from those born in [1930–39] to those born in [1970–1974]) are defined.

These groups are refined further by incorporating three educational groupings: [1] some or completed primary school, [2] some or completed secondary school, and [3] some university or university degree. Thus, each individual observation is classified into one of 93 cohort-age-education groups as shown in Table 5, which also displays the number of observations per cell. Only nine cells have fewer than two hundred observations, meeting the criterion cited by Verbeek and Nijman (1992; 1993) for nearly unbiased estimates from pseudo-panel procedures.

In examining the relationship between working hours and wages, group means are constructed for the sample of working women only. The principal variables used in the estimation are weekly working hours and weekly wages. The weekly working hours variable has been directly drawn from the survey, while the weekly wages come from division of the reported monthly wage by a constant (4.3).[2] While hourly wages would be the ideal measure of the opportunity cost of labour, earnings are reported as hourly wages by only one per cent of wage earners (Nakavachara 2010). Although hourly wages can be computed from reported weekly earnings, this approach can lead to a problem of 'division bias' when errors in reported hours contaminate both right and left hand side variables (Borjas 1980; Welch 1997).[3] Reported wages are deflated by the Thailand Consumer Price Index (CPI).[4]

Additional controls incorporated in specific models include a business cycle measure and demographic characteristics. The former is captured by the annual rate of change in real GDP associated with the survey year of each cell. Demographic factors include marital status (proportions married and non-married) and proportions living in urban versus rural areas. These demographic variables are incorporated first as controls in alternative specifications of the aggregate female labour supply relations, and then to define disaggregated samples to investigate differences in labour supply equations across demographic groups.

The basic summary statistics in Table 1 show that women earned on average 2258THB per week; while their working hours averaged around 43 hours per week. The employment population ratio states that 83 per cent of women in our sample were employed. Workers in our sample were more likely to obtain university degrees (50 per cent), live in urban area (70 per cent), and be married (60 per cent). In addition, Tables 2–4 present average real weekly earnings, average weekly hours, and the employment-population ratio disaggregated by age, birth cohort, and educational attainment.

In Table 2, average real weekly wages increase monotonically with age for all three education groups, and this is true even controlling for birth cohort (reading horizontally across any row). This age profile of earnings may reflect increased productivity as women gain more experience in the workforce. However, surprisingly, wages have declined steadily by birth cohort for all three education groups, but this is true only when not controlling for age. Within each age group, the wage pattern for the different birth cohorts is mixed, pointing to the need for regression analysis with suitable controls.

These data show either declining or mixed trends in wages which seem to contradict other evidence on the strong upward trends in average wages observed through 1995 (Nakavachara 2010, figure 2). The subsequent stagnation of wages through 2005 has

[2]The survey does not report information on weeks worked per month, which we treat as a constant.

[3]Concern over possible division bias is reinforced by estimates of the hours worked equation using the constructed hourly wage data from the LFS. The estimated elasticities with respect to hourly wages using pseudo-panel methods are similar to those based on individual data (estimates not reported here but available on request). Since regressions using individual data show evidence of inconsistency, both from theoretical considerations and in comparison with the pseudo-panel estimates using reported weekly earnings, the estimated elasticities for the hourly wage data are apparently biased as well.

[4]The CPI indexes (2002 as a base year) are from the Bureau of Trade and Economic Indices, Ministry of Commerce, Thailand.

Table 1. Summary statistics of full sample (individual level).

	(1)	(2)	(3)	(4)	(5)	(6)
Variables	Unit	Observations	Mean	Standard deviation	Min	Max
Age	Year	109,095	36.33	8.168	25	60
Real GDP	%	109,095	4.981	5.432	−10.50	13.20
Weekly Wage	Baht	109,095	2258	1669	23.26	40,665
Weekly Working Hours	Hour	109,095	43.45	11.57	1	98
Employment-Population Ratio	%	109,095	0.828	0.079	0.528	0.956
Working Hours-Population Ratio	%	109,095	35.70	8.820	0.681	92.81
Education					[95% confidence Interval]	
Proportion of Women with Primary	%	109,095	0.260	0.439**	0.257*	0.262*
Proportion of Women with Secondary	%	109,095	0.232	0.422**	0.229*	0.234*
Proportion of Women with University	%	109,095	0.507	0.500**	0.504*	0.510*
Living Area					[95% confidence Interval]	
Proportion of Women Living in Rural	%	109,095	0.299	0.458**	0.296*	0.302*
Proportion of Women Living in Urban	%	109,095	0.700	0.458**	0.697*	0.703*
Marital Status					[95% confidence Interval]	
Proportion of Non-Married Women	%	109,095	0.391	0.488**	0.389*	0.394*
Proportion of Married Women	%	109,095	0.608	0.488**	0.605*	0.610*

*95% confidence interval for proportions are calculated using a logit transform.
**Standard deviations are obtained from standard errors and confidence intervals for group means.

Table 2. Average real weekly wage by cohort and age.

	Age Group							
Cohort	25–30	31–35	36–40	41–45	46–50	51–55	56–60	All
Education = 1 (primary level)								
Born 1930–1939					986	1187	1369	1221
Born 1940–1944				917	1004	1297	1753	1268
Born 1945–1949			869	1057	1144	1599	2159	1318
Born 1950–1954		893	941	1157	1268	1594		1162
Born 1955–1959	706	860	1056	1177	1388			1050
Born 1960–1964	740	1005	1118	1267				1019
Born 1965–1969	951	1101	1127					1042
Born 1970–1974	1101	1198						1127
All	939	1039	1062	1163	1223	1458	1707	1104
Education = 2 (secondary level)								
Born 1930–1939					2317	3040	3980	3301
Born 1940–1944				1930	2799	3684	4044	3200
Born 1945–1949			2012	2634	3428	3947	4752	3130
Born 1950–1954		1616	2175	2733	3422	4769		2619
Born 1955–1959	1273	1732	2263	2530	3249			2055
Born 1960–1964	1265	1736	1848	2390				1660
Born 1965–1969	1473	1615	1765					1549
Born 1970–1974	1461	1603						1484
All	1411	1670	2018	2531	3219	3764	4086	1945
Education = 3 (university level)								
Born 1930–1939					3495	3998	5070	4394
Born 1940–1944				2730	3747	4816	5358	4420
Born 1945–1949			2565	3495	4577	4939	5967	4182
Born 1950–1954		2164	3083	4057	4545	5592		3761
Born 1955–1959	1748	2414	3293	3754	4662			3117
Born 1960–1964	1822	2691	2993	3676				2660
Born 1965–1969	2242	2593	3129					2483
Born 1970–1974	2129	2623						2222
All	2068	2567	3018	3780	4445	4833	5039	2991

occurred despite continued educational gains at the upper secondary and university levels (Nakavachara 2010, Table 3). Since there are large wage differentials by level of education, these continued gains in female educational attainment have been offset by wage declines within each educational group (Table 2).

The female employment ratio increases by year of birth for all three education categories (Table 4, last column). However, after controlling for age, the trends in the employment ratio are mixed. Those with some university education show declines by year of birth for all age groups except the youngest (25–30 year olds). The mixed patterns of wages and

Table 3. Average weekly hours worked by cohort and age.

	Age Group							
Cohort	25–30	31–35	36–40	41–45	46–50	51–55	56–60	All
Education = 1 (primary level)								
Born 1930–1939					49	48	49	49
Born 1940–1944				51	51	49	47	49
Born 1945–1949			51	51	51	48	44	49
Born 1950–1954		52	52	51	49	48		51
Born 1955–1959	54	54	52	51	49			52
Born 1960–1964	56	53	51	51				53
Born 1965–1969	55	52	51					53
Born 1970–1974	54	51						53
All	55	53	52	51	50	48	47	52
Education = 2 (secondary level)								
Born 1930–1939					38	38	37	37
Born 1940–1944				39	39	38	36	38
Born 1945–1949			40	41	39	38	38	39
Born 1950–1954		41	41	40	40	40		41
Born 1955–1959	43	43	41	42	40			42
Born 1960–1964	45	44	44	43				44
Born 1965–1969	47	46	45					47
Born 1970–1974	48	47						48
All	47	45	43	41	39	38	37	44
Education = 3 (university level)								
Born 1930–1939					37	37	36	36
Born 1940–1944				38	37	36	36	36
Born 1945–1949			37	37	37	37	39	37
Born 1950–1954		37	38	36	37	38		37
Born 1955–1959	38	38	37	37	39			37
Born 1960–1964	41	39	39	40				39
Born 1965–1969	41	40	42					41
Born 1970–1974	41	42						42
All	41	39	38	37	37	37	36	39

Table 4. Employment-population ratio.

	Age Group							
Cohort	25–30	31–35	36–40	41–45	46–50	51–55	56–60	All
Education = 1 (primary level)								
Born 1930–1939	0	0	0	0	0.749	0.691	0.584	0.644
Born 1940–1944	0	0	0	0.780	0.768	0.680	0.575	0.674
Born 1945–1949	0	0	0.783	0.803	0.753	0.681	0.637	0.729
Born 1950–1954	0	0.770	0.812	0.792	0.753	0.714	0	0.771
Born 1955–1959	0.742	0.800	0.813	0.799	0.784	0	0	0.794
Born 1960–1964	0.763	0.789	0.805	0.818	0	0	0	0.793
Born 1965–1969	0.758	0.793	0.832	0	0	0	0	0.787
Born 1970–1974	0.748	0.799	0	0	0	0	0	0.762
All	0.754	0.791	0.810	0.799	0.760	0.689	0.590	0.757
Education = 2 (secondary level)								
Born 1930–1939	0	0	0	0	0.640	0.616	0.528	0.578
Born 1940–1944	0	0	0	0.703	0.716	0.663	0.544	0.656
Born 1945–1949	0	0	0.721	0.727	0.722	0.671	0.553	0.705
Born 1950–1954	0	0.726	0.769	0.771	0.754	0.684	0	0.758
Born 1955–1959	0.694	0.753	0.767	0.794	0.760	0	0	0.753
Born 1960–1964	0.714	0.762	0.779	0.772	0	0	0	0.754
Born 1965–1969	0.746	0.768	0.813	0	0	0	0	0.761
Born 1970–1974	0.751	0.783	0	0	0	0	0	0.756
All	0.738	0.763	0.775	0.764	0.732	0.660	0.538	0.746
Education = 3 (university level)								
Born 1930–1939	0	0	0	0	0.932	0.869	0.776	0.833
Born 1940–1944	0	0	0	0.935	0.926	0.932	0.754	0.88
Born 1945–1949	0	0	0.956	0.931	0.947	0.877	0.686	0.913
Born 1950–1954	0	0.930	0.929	0.938	0.921	0.767	0	0.922
Born 1955–1959	0.867	0.914	0.933	0.927	0.859	0	0	0.915
Born 1960–1964	0.849	0.893	0.901	0.869	0	0	0	0.883
Born 1965–1969	0.886	0.881	0.868	0	0	0	0	0.881
Born 1970–1974	0.875	0.87	0	0	0	0	0	0.874
All	0.872	0.892	0.912	0.921	0.919	0.876	0.754	0.893

the employment ratio across cohorts and age groups calls for a more systematic analysis of the relation between these two variables.

With female employment at such high levels, greater involvement in work must come from increases in weekly hours (Table 3). For all three education categories, there is a tendency for hours worked to decline with age. On the other hand, hours worked have generally increased with year of birth, although only weakly for those in the lowest education category.

The trends and patterns in the three key indicators of female labour markets are subtle and complex. Aggregate trends may disguise the underlying relations between female labour supply and its determinants, requiring formal data analysis with appropriate controls. In particular, trends and patterns in the three variables depend on whether or not one controls for age, birth cohort, or educational attainment. The use of synthetic cohorts in the regression analysis allows such controls, in order to capture the effect of wages on hours worked that is purged of the influence of human capital accumulation through prior labour market activities.

IV. Empirical strategy for labour supply estimation

Studies of female labour supply in developing countries using individual cross-sectional data produce biased estimates due to omitted individual effects. Panel data analysis allows for control of unobserved individual effects, but such data sets are seldom available in developing countries. Statistical bureaus in many developing countries collect cross sectional survey data over time, sampling a different set of households or individuals in each survey year. Since different individuals are sampled each year, this sequence of cross sectional surveys is not a true panel, and standard panel estimators such as least squares – dummy variables (LSDV) cannot be applied directly to solve problems of individual heterogeneity bias.

In the absence of panel data, Deaton (1985) proposes the construction of pseudo-panel data from the time series of cross-sectional surveys. For each year, the individual observations are averaged within groups, for example, within birth cohorts. These group means become the cross sectional units of observation, and these group means are observed over time as in a panel data set. Dummy variables representing each group can be included in the panel estimation to control for differences across groups, as in conventional fixed effects estimation, in order to reduce the problem of heterogeneity bias. Angrist and Pischke (2009) demonstrate that generalized least squares estimation applied to group means is equivalent to instrumental variables estimation with individual data where dummy variables indicating group membership serve as instruments. In addition to coping with the problem of heterogeneity bias, Antman and McKenzie (2007) point out that application of pseudo-panels deals with three problems common to developing country data sets: the absence of sufficiently long panels; attrition of individuals from longitudinal surveys; and substantial measurement errors characteristic of individual survey responses. With these advantages of pseudo-panels, this methodology is beginning to see applications in development economics research, such as Contreras, Bravo, and Puenes (2005), Antman and McKenzie (2007), Warunsiri and McNown (2010), Paweenawat and McNown (2014), and Motonishi (2006).

The fundamental equation in this study is the relation between the quantity of labour supplied (log of weekly hours or the employment/population ratio – $\ln h_{it}$) and the log of the weekly wage rate ($\ln w_{it}$), augmented by a vector of appropriate control variables (X_{it}), each observed for individual women, indexed by i, at different survey years, indexed by t:

$$\ln h_{it} = \gamma + \theta \ln w_{it} + X_{it}'\eta + \alpha_i + u_{it} \qquad (1)$$

All variables in (1) are observed for employed women in the equation for hours worked; for the employment/population ratio equation both working and non-working women are included in the sample. Unobservable individual heterogeneity, such as motivation or ability, is represented by α_i. In the likely case that α_i is correlated with wages, least squares estimators of this relation will be biased and inconsistent. Following Pencavel (1998), individual observations are aggregated into groups defined by age at the time of the survey (indexed by *a*), year of birth (indexed by *k*), and level of educational attainment (index *j*), and all variables in (1) are averaged over those women within each of these groups.

The labour supply equation in terms of group means becomes

$$\overline{\ln h_{akj}} = \gamma + \theta \overline{\ln w_{akj}} + \overline{X_{akj}}'\eta + \overline{\alpha_{akj}} + \overline{u_{akj}}, \qquad (2)$$

where $\overline{\ln h_{akj}}$ is the average of the log of working hours per week of age group a in birth cohort k with education level j, and similarly for the other terms in this equation. For example, the first non-zero observation on wages in the pseudo panel is the average wage for working women who were aged 46–50 at the time of the survey, born between 1930 and 1939, and having no more than a primary school education (Table 5). All three subscripts are necessary to identify the categories of women defining the pseudo-panel observations. A total of 93 non-zero group means are constructed from the 109,905 individual observations (Table 5).

Construction of pseudo-panel data sets involves a trade-off between estimator efficiency and minimal bias. Pseudo-panel estimators are biased if cell sample sizes are too small (Deaton 1985). On the other hand, cell sample sizes can be increased by defining broader categories of individuals for construction of group means, e.g. ten-year birth cohorts versus five-year birth cohorts. Such aggregation defines fewer groups and therefore fewer observations in the pseudo-panel estimation, and larger standard errors for the parameter estimates. Verbeek and Nijman (1992, 1993) find that cell sample sizes greater than 100 observations are generally sufficient to produce estimators with small biases.

In addition, heterogeneity across cohorts can produce biased pseudo-panel estimators in the absence of group-specific fixed effects. Estimates with and without various fixed effects, as well as other control variables, are presented to show how alternative sources of variation affect elasticity estimates. The control variables differ across specifications and include:

(1) Age effects (f_a or age dummies),
(2) Birth cohort effects (f_c or cohort dummies),
(3) Education effects (or the education level dummies)
(4) The business cycle effect (or the average annual change of real GDP for that group)
(5) Demographic factors (marital status and living areas, both as group proportions)

The key parameter of interest is θ, the coefficient on the logarithm of weekly wages, which is the wage

Table 5. Number of observations per cohort-age-education cell.

	Age Group							
Cohort	25–30	31–35	36–40	41–45	46–50	51–55	56–60	Total
Education = 1 (primary level)								
Born 1930–1939	0	0	0	0	189	333	376	898
Born 1940–1944	0	0	0	243	318	374	327	1262
Born 1945–1949	0	0	399	431	567	645	248	2290
Born 1950–1954	0	505	748	964	1096	438	0	3751
Born 1955–1959	815	831	1463	1629	664	0	0	5402
Born 1960–1964	1389	1545	1978	860	0	0	0	5772
Born 1965–1969	2283	1970	1066	0	0	0	0	5319
Born 1970–1974	2708	973	0	0	0	0	0	3681
Total	7195	5824	5654	4127	2834	1790	951	28,375
Education = 2 (secondary level)								
Born 1930–1939	0	0	0	0	116	168	233	517
Born 1940–1944	0	0	0	176	282	370	187	1015
Born 1945–1949	0	0	333	368	552	380	49	1682
Born 1950–1954	0	403	557	641	503	81	0	2185
Born 1955–1959	847	907	1304	931	203	0	0	4192
Born 1960–1964	1845	1798	1508	425	0	0	0	5576
Born 1965–1969	3308	1913	579	0	0	0	0	5800
Born 1970–1974	3662	690	0	0	0	0	0	4352
Total	9662	5711	4281	2541	1656	999	469	25,319
Education = 3 (university level)								
Born 1930–1939	0	0	0	0	188	306	429	923
Born 1940–1944	0	0	0	267	413	740	465	1885
Born 1945–1949	0	0	556	727	1309	881	121	3594
Born 1950–1954	0	954	1486	2419	1781	229	0	6869
Born 1955–1959	1587	2194	4364	3351	529	0	0	12,025
Born 1960–1964	2948	4290	4256	906	0	0	0	12,400
Born 1965–1969	4939	3987	1171	0	0	0	0	10,097
Born 1970–1974	6178	1430	0	0	0	0	0	7608
Total	15,652	12,855	11,833	7670	4220	2156	1015	55,401

elasticity of labour supply for the hours-worked equation. In addition, to investigate the extensive margin of labour supply, the employment-population ratio (E_P) and working hours-population ratio (H_P) are used as alternative dependent variables. For each of the 93 groups of women defined by age, birth cohort, and education level, E_P is computed as the proportion of women who are employed, and the logarithm of this ratio becomes the dependent variable in a new set of regressions. Similarly, the average number of hours worked for all women in each of these same 93 groups defines a third alternative dependent variable, H_P, also used in logarithmic form. The latter measure combines both intensive and extensive margins to show an overall labour supply response to wage changes.

Since the hours worked equation is estimated with a sample based on working women only, it is necessary to consider potential selection bias problems. Studies of labour supply using grouped data by Blundell, Duncan, and Meghir (1998) deal with this problem by including the inverse Mills ratio, constructed from the proportion of women in each group who are employed. We adopt the same strategy for some specifications, controlling for employment status and/or marital status when appropriate.

Each equation specification is estimated by both weighted least squares and instrumental variables procedures, the latter accommodating wage endogeneity. Feedback from hours worked to wages is possible if labour markets are segmented by women's age, birth year, and education level. IV estimation can also treat other sources of wage endogeneity, such as wage misreporting.

Following the recommendation of Angrist and Pischke (2009), we check the sensitivity of estimates to alternative identifying assumptions, in particular, allowing cohort wages to be endogenous. An appropriate instrument for cohort wages is cohort educational attainment, based on the evidence of returns to education studies (e.g. Warunsiri and McNown 2010 for Thailand, in particular). Pencavel (1998), Blundell, Duncan, and Meghir (1998), and Schultz (1990) use education as an instrument in their female labour supply studies, in some cases combined with other exogenous variables. In order to use educational attainment as an instrument, it must be excluded as a control variable in the labour supply equation. In this case the estimated response of labour supply to wage changes incorporates the effect that future wage expectations have on schooling decisions. Other studies, such as Pencavel (1998), have chosen to exclude education as a control in order to estimate this more comprehensive effect of wages on the supply of labour. The choice of education proportions as an appropriate instrument is subjected to tests for weak instruments. In this study, estimates from alternative models with a variety of controls are presented to show the sensitivity of results to these alternative specifications.

The final econometric issue to be addressed is heteroscedasticity arising from the different sample sizes used to compute the cohort means. Weighted least squares (WLS) estimates are reported in the table of estimates, where the weights are equal to the square root of the number of observations in each cell (Dargay 2007).

V. Estimates of female labour supply relations for Thailand

Age patterns in wages, working hours, and employment

Overall patterns for the three key indicators of labour market activity were summarized previously, noting that these patterns differ depending on birth cohort, age, or educational attainment controls. The goals of the econometric analysis are to pin down the elasticities of labour supply for both intensive and extensive margins with respect to wages, and to investigate how elasticities vary under different controls and within alternative demographic groups. Disaggregating the sample by birth cohort, marital status, educational attainment, or urban vs. rural residence, differences in elasticity estimates shed light on the evolving patterns and trends in female labour markets in Thailand.

Female labour supply relations for all women

Table 6 displays estimates of responses of Thai women's labour market activity to changes in wages. The first column presents the pooled regression estimates based on individual data, and the next two columns present WLS and IV estimates of the elasticity of weekly hours worked with respect to weekly earnings for all working women. As discussed in the methodology Section IV, the

proportion of women of each cohort in the three educational categories has been employed as the instrumental variable.

A comparison of the WLS and IV estimates with those in the first column shows the reduction in biases from controlling for individual heterogeneity. The fourth column reports WLS estimates of the response of the employment/population ratio (E_P) to changes in the log of wages, while the last column presents the WLS estimates of the response of the log of working hours-population ratio (H_P) to changes in the log of wages. Estimates are reported for six alternative specifications defined by different selections of control variables.

Generally, the labour supply estimates are quite robust across the alternative specifications when estimated with the pseudo-panel methods. However, the estimated hours worked elasticity based on individual data is sharply different from all pseudo-panel estimates; (approximately −0.13 compared with a range of −0.22 to −0.25 using the pseudo-panel data). This comparison provides convincing evidence of heterogeneity bias in labour supply estimates from using data on individuals, which has been the case for most prior studies in developing countries. A likely candidate for this unobserved heterogeneity is motivation or ability, as is commonly assumed in models of returns to education (see Warunsiri and McNown 2010 for a discussion of this in the context of Thailand). If this omitted ability factor is positively correlated with both wages and hours worked, the positive heterogeneity bias found here is consistent with this interpretation. Hours worked by Thai women, in aggregate, are twice as responsive to wage changes as would be predicted by individual OLS regressions. From Dessing's (2002) interpretation this estimated elasticity implies a strong desire to reduce work hours in favour of other household activities, when an increase in women's wages permits.

Focusing on the pseudo-panel regressions, the relation between hours worked and wages is negative, with elasticity estimates falling within the tight range of −0.22 and −0.25. Overall, the coefficients are similar for WLS and IV estimation, with neither estimate consistently smaller or larger than the other. The largest difference between WLS and IV estimated elasticities is 0.03. Such small differences indicate that endogeneity in the pseudo-panel estimates is not a concern for these labour market data. To reinforce this conclusion, evidence on the validity of the educational category instruments is provided by estimates and tests of the first-stage regressions. In most cases, the educational category variable

Table 6. Estimates of the association between work and wage.

		Pseudo-panel			
Controls	Cross-sectional Individual	WLS	IV	E_P	H_P
1. None	−0.132*	−0.223*	−0.223*	0.937*	−0.116*
	(0.0013)	(0.0086)	(0.0254)	(0.0110)	(0.0159)
2. Age and cohort	−0.130*	−0.234*	−0.223*	0.110*	−0.108*
	(0.0015)	(0.0079)	(0.0129)	(0.0083)	(0.0134)
3. Age, cohort, and selected adjusted		−0.246*	−0.252*		
		(0.0165)	(0.0267)		
4. Age, cohort, and business cycle	−0.129*	−0.233*	−0.223*	0.106*	−0.0756*
	(0.0015)	(0.0079)	(0.0129)	(0.0081)	(0.0156)
5. Age, cohort and demographic					
5.1 Married/non-married	−0.128*	−0.227*	−0.223*	0.110*	−0.0789*
	(0.0015)	(0.0065)	(0.0122)	(0.0056)	(0.0113)
5.2 Urban/rural	−0.131*	−0.233*	−0.223*	0.110*	−0.0775*
	(0.0015)	(0.0061)	(0.0121)	(0.0062)	(0.0121)
Number of observations	109,095	93	93	93	93

Heteroscedastic-consistent estimated standard errors are in parentheses. WLS stands for weighted least-squares.
IV stands for instrumental variable, which is the proportion of women of each cohort in three educational categories.

- Line 1 – Weak instruments test on first stage regression: Robust F (1,91) = 26.92 (p = 0.00)
- Line 2 – Weak instruments test on first stage regression: Robust F (1,78) = 50.70 (p = 0.00)
- Line 3 – Weak instruments test on first stage regression: Robust F (1,77) = 14.65 (p = 0.00)
- Line 4 – Weak instruments test on first stage regression: Robust F (1,77) = 24.91 (p = 0.00)
- Line 5.1 – Weak instruments test on first stage regression: Robust F (1,77) = 43.93 (p = 0.00)
- Line 5.2 – Weak instruments test on first stage regression: Robust F (1,77) = 32.84 (p = 0.00)

The coefficient on the inverse Mills ratio and its standard error in WLS are −47.16 and 38.56 (p = 0.22), respectively.
The coefficient on the inverse Mills ratio and its standard error in IV are −0.34 and 0.41 (p = 0.40), respectively.
*Significant at 0.05 significant level.

passes the test for weak instruments, as shown in Tables 6–9. The equations are exactly identified with this single IV, and therefore no test for over-identifying restrictions applies.

Comparing line 1 estimates with those of other specifications shows how different controls affect estimated elasticities. Inclusion of age and cohort dummies, for example, produces estimates that use only variation across education groups and over time. Addition of these controls increases the WLS estimate of the absolute value of the wage elasticity from 0.223 to 0.234, and the IV estimates are identical to three decimal places (line 2 vs. line 1).

Other model changes result in only small differences from the line 2 specification in terms of the estimated elasticity of hours worked with respect to earnings. In particular, the correction for sample selection in line 3 slightly raises the absolute elasticity for the WLS estimate (from −0.223 to −0.246) and the IV estimate (from −0.223 to −0.252). The lack of statistical significance on the coefficient of inverse Mills ratio is consistent with the finding that the IV estimates differ little from the WLS estimates. The coefficient on the inverse Mills ratio and its standard errors have been reported at the end of Table 6. It appears that sample selection is not an important issue in this labour supply equation – a result that squares with Pencavel's (1998) finding for U.S. data. In addition, neither business cycle (line 4), nor demographic (line 5) controls produce much change in the estimates of the basic wage elasticity.

In every case, there is a clear finding of a negatively sloped supply curve, a result that is consistent with other studies of female labour supply in developing countries. In a study using Thai survey data for 1980–81, Schultz (1990) finds a negative and statistically significant coefficient on wages in his hours worked equation for women. Schultz does not report an elasticity estimate. Similarly, Dasgupta and Goldar (2006) report a negative and significant coefficient on wages in their study of labour supply of Indian women below the poverty line, again with no elasticity estimate given.

Other studies confirm the inelasticity of supply reported here. Yamada (2008) uses micro data in cross sectional estimations of labour supply functions for low income individuals in Lima, Peru. His estimated elasticity for women is −0.13, which matches our estimate based on individual data (column 1 of Table 6). Dessing (2002) presents an empirical analysis based on a 99 person 1976 survey of low-income individuals in rural Philippines. Although the small sample size weakens the statistical significance of her estimates, wage elasticity estimates for secondary workers range from −0.16 to −0.46 depending on estimation methods and control variables.

Several studies consider the possibility that the labour supply function could be positively sloped at higher income levels, while negatively sloped supply curves are characteristic of secondary workers, especially at lower levels of income. Sharif (1991) uses Indian data from 1976 to estimate labour supply functions that allow both rising and falling sections. These functions are estimated for twelve different demographic groups, and half of these groups have both positive and negative sloped sections of the supply curve. The one for adult women 15–44 years of age, which comes closest to this current study, exhibits only a negatively sloped supply function. At the average wage for this group, the estimated elasticity is −0.26. In a similar exercise with the Thai pseudo-panel data of this study, the hours worked equation was estimated with the sample segmented into three different levels of wages.[5] The result still shows a downward sloping female labour supply for all three levels of wages. However, workers in the highest wage group have the lowest absolute wage elasticity (−0.105) of the three, so that the labour supply curve could be approaching a flat or even a positive slope at higher levels of income.

Estimates of female labour supply that incorporate labour force participation are reported in columns 4 and 5, where the employment-population ratio (E_P) and working hours-population ratio (H_P) are used as alternative dependent variables. As in the hours-worked equation, these two dependent variables are also constructed as mean values for each age, birth cohort, and education group. Since WLS and IV estimates of female labour supply relations showed little difference, only WLS

[5]Estimates of work-wage association for the bottom-third of income, second-third of income, and top-third of income are −0.220, −0.287, and −0.105, respectively.

estimates are reported. The employment/population ratio responds positively to a wage increase, with a 10 per cent increase in real wages raising the employment/population ratio by approximately 1 percentage point, an estimated effect that is similar across specifications. For all specifications with age and cohort controls, wage coefficients are all equal (0.11) to the second decimal place.

Combining both intensive and extensive margins into an average of hours-worked per woman (H_P), the net effect of a rise in wages on the quantity of labour supplied is still negative. The range of estimates is greater than those for the intensive and extensive margins taken separately. A 10 per cent increase in women's wages lowers the working hours-population ratio between 0.76 and 1.16 percentage points across the various specifications. The positive effect of wages on the employment ratio is small because of the initial high level of labour market participation among Thai women; there is simply not much scope for increasing the number of employed women through increased wages. This relatively small positive effect from the extensive margin is more than offset by the negative impact on the intensive margin, reinforcing the weakness of the substitution effect compared to the income effect in the labour supply behaviour of women in low income households.

The estimates of female labour supply based on the complete sample are quite robust across a variety of model specifications and estimation procedures. Not only is the hours worked elasticity with respect to wages consistently negative, but the range of estimates is quite narrow. Estimates of the response of the employment/population ratio are also tightly clustered. This uniformity of estimates based on the complete sample hides some important differences in female labour supply responses for different demographic groups, as discussed in the following section.

Disaggregated estimates of the association between work and wages

In this section, the data are disaggregated to allow estimates to differ across demographic groups (birth cohort, age group, marital status, urban vs. rural residence, and education level). Since the estimated elasticities for all women do not change substantially with different controls (Table 6), the disaggregated equations include a minimum of controls to allow variation across multiple dimensions to enter into the estimation. In addition, when separate models are estimated for each demographic group, models that include these same demographic variables as controls are not identified. Disaggregation by demographic characteristics is a more informative approach to controlling for these factors since elasticities are allowed to differ across these groups.

Disaggregation by cohort and by age group

Labour supply relations are estimated separately for three birth cohorts using an equation with no controls (Table 7). The most recent birth cohorts exhibit the greatest response to wage changes, in terms of both increases in employment probabilities and decreases in hours worked resulting from a rise in women's wages. Since the pattern across age groups is the same, with the youngest women showing the greatest responsiveness to wage changes, a separate table of estimates with the data disaggregated into three age groups is not shown here.

Older women and those from earlier birth cohorts are likely to have established lifestyle and work patterns and may be more strongly governed by traditional mores than younger women and those born more recently. They would, therefore, be less responsive to wage changes in decisions over whether or not to work, as reflected in smaller wage coefficients in the E_P equations. These coefficients rise monotonically with year of birth over the three birth cohorts, implying that a higher employment elasticity with respect to wages could emerge in the future.

Among women who are already employed, the estimates of Table 7 indicate that younger workers from more recent birth cohorts are more responsive to wage changes in their decisions over working hours. There are several possible explanations of this result, including differences in health or vitality and family obligations and sources of support. First, in response to a wage reduction, younger, healthier women are better able to increase their working hours, especially in those occupations requiring physical strength. Second, in Thai culture people tend to live together with members of their extended family. Older female workers might not be the primary wage earners in the extended household, so that their wage changes may have little impact on total household income. Consequently, their number of working hours is less responsive to a given wage change, compared with the primary workers of the household.

Table 7. Estimations of the association between work and wage by cohort.

Cohort	Number of Observations	WLS	IV	E_P
Born 1930–1949	36	−0.182*	−0.140*	0.0814*
		(0.0071)	(0.0418)	(0.0271)
Born 1950–1959	30	−0.209*	−0.228*	0.0979*
		(0.0124)	(0.0464)	(0.0128)
Born 1960–1974	27	−0.273*	−0.263*	0.122*
		(0.0161)	(0.0267)	(0.0953)
All	93	−0.223*	−0.223*	0.0937*
		(0.0868)	(0.0254)	(0.0110)

Heteroscedastic-consistent estimated standard errors are in parentheses. WLS stands for weighted least-squares.
IV stands for instrumental variable, which is the proportion of women of each cohort in three educational categories.

- Born 1930–1949 – Weak instruments test on first stage regression: Robust F (1,34) = 3.74 (p = 0.06)
- Born 1950–1959 – Weak instruments test on first stage regression: Robust F (1,28) = 10.44 (p = 0.00)
- Born 1960–1974 – Weak instruments test on first stage regression: Robust F (1,25) = 47.72 (p = 0.00)
- All – Weak instruments test on first stage regression: Robust F (1,91) = 26.92 (p = 0.00)

*Significant at 0.05 significant level.

In addition, these older women might have greater family burdens, such as housework or taking care of children, than younger workers, so that they are less able to adjust their working hours due to other family responsibilities.

Disaggregation by marital status

Table 8 provides evidence on how labour supply relations differ by marital status. For women at all ages, the response of the employment ratio to wage changes differs very little between married and unmarried women (columns headed by E_P). When the data are disaggregated by the three age groups, these responses are smaller for unmarried women. This finding is contrary to the predictions of the trailing spouse hypothesis, which contends that the labour supply behaviour of married women is more constrained by their marital and family commitments. However, the estimates reported in Table 8 indicate that married women have more flexibility in their choice of whether or not to work because of the cushion of their husband's income. This is similar to Mammen and Paxson's (2000) finding that higher husband's earnings of younger cohorts provides greater flexibility in the woman's decision to work or not. On the other hand unmarried women, who lack a husband's income for support, may be unable to withdraw from the labour market even as wages become less attractive.

In the hours worked equations, unmarried women have consistently greater absolute wage elasticities than their married counterparts. The interpretation here is similar to the discussion of the extensive margin. Facing a decline in wages, unmarried women must increase their hours worked to maintain basic expenditure levels. Married women facing a similar wage shock, on the other hand, can rely on spouse's earnings and will not increase their own hours of work as dramatically.

Disaggregation by place of residence

Labour supply responses for women living in urban areas and rural areas are shown in Table 9 for three age categories. In the hours worked equations for women of all ages, the estimated elasticities are similar for urban and rural women. When broken into the three age groups, comparisons of the urban and rural elasticities show no consistent pattern. Only in the equations for the employment rate are the estimates consistent with the hypothesis of gender differences in the urban labour market. For each age group and for women of all ages, the rural employment rate response to wage changes is greater than that in the cities. It is often argued that the growth of urban jobs during early stages of industrialization favour men and are not suitable for women. This view of early industrialization is consistent with the finding that urban female labour supply is less responsive to wage changes than for women living in rural areas. This result also reinforces Mammen and Paxson (2000) observation of higher female labour force participation rates in rural compared with urban areas.

Disaggregation by educational attainment

Thailand has promoted schooling of girls as part of its goals of gender equality, and these policies have led to female education levels now matching or even exceeding those for males (Knodel 1997; Nakavachara 2010). With an increasing

Table 8. The relationship between work and wage for non-married/married women by age.

	Married			Non-married		
Age Group	WLS	IV	E_P	WLS	IV	E_P
25–35	−0.261*	−0.290*	0.140*	−0.302*	−0.324*	0.125*
	(0.0196)	(0.0406)	(0.0169)	(0.0218)	(0.0386)	(0.0129)
36–45	−0.217*	−0.204*	0.0987*	−0.267*	−0.247*	0.0935*
	(0.00948)	(0.0211)	(0.0113)	(0.0117)	(0.0219)	(0.0096)
46–60	−0.158*	−0.109*	0.0995*	−0.199*	−0.137*	0.0874*
	(0.00665)	(0.0333)	(0.0261)	(0.00715)	(0.037)	(0.0244)
All	−0.201*	−0.210*	0.0933*	−0.259*	−0.243*	0.0940*
	(0.00807)	(0.0266)	(0.0127)	(0.0119)	(0.0249)	(0.0109)

Heteroscedastic-consistent estimated standard errors are in parentheses. WLS stands for weighted least-squares.
IV stands for instrumental variable, which is the proportion of women of each cohort in three educational categories.

- Married – Weak instruments test on first stage regression: Robust F (1,91) = 25.00 (p = 0.00)
  - Age group (25–35) – Weak instruments test on first stage regression: Robust F (1,25) = 33.20 (p = 0.00)
  - Age group (36–45) – Weak instruments test on first stage regression: Robust F (1,28) = 24.24 (p = 0.00)
  - Age group (46–60) – Weak instruments test on first stage regression: Robust F (1,34) = 5.05 (p = 0.03)
- Non-Married – Weak instruments test on first stage regression: Robust F (1,91) = 31.37 (p = 0.00)
  - Age group (25–35) – Weak instruments test on first stage regression: Robust F (1,25) = 35.36 (p = 0.00)
  - Age group (36–45) – Weak instruments test on first stage regression: Robust F (1,28) = 30.62 (p = 0.00)
  - Age group (46–60) – Weak instruments test on first stage regression: Robust F (1,34) = 5.58 (p = 0.02)

*Significant at 0.05 significant level.

proportion of female workers pursuing secondary and tertiary schooling, how are labour supply relations changing? Estimation of labour supply equations with the data disaggregated according to three levels of educational attainment provides a perspective on this question.

Table 10 shows that the absolute wage elasticities of hours worked for women with some university education are lower than those with primary or secondary schooling only.[6] Assuming highly educated women work in the most attractive jobs, the amenities associated with these jobs weaken the response of women in these positions to wage changes. On the extensive margin, women with some primary school education are not responsive to wage changes in the decision to work, even though their overall attachment to the labour force is at least as strong as those with some secondary schooling. Women with secondary schooling respond positively to an increase in wages in their employment decision, and the employment ratio for those with some university education is the most responsive to wages. As women's educational attainment continues to rise, so does their attachment to employment, as well as the responsiveness of the employment ratio to changing wages. One implication is that continued increase in women's educational attainment and rising wages will both contribute to the supply of the most productive women to the labour market.

VI. Conclusion

Over the two decades after 1985, the Thai economy produced substantial growth, industrialization, and modernization, accompanied by corresponding changes in the labour force. The proportion of wage and salary workers grew from 27 per cent in 1985 to 48 per cent in 2005.[7] This formal sector workforce is more highly educated (25 per cent with tertiary education in 2004 vs. 12 per cent in 1985), increasingly feminized (45 per cent of wage and salary workers in 2004 were women), and better paid (average earnings have increased by 60 per cent over this period). Increases in girls' education and shifting attitudes towards gender equality have reduced the earnings gap between men and women to 9 per cent by the end of this period. Some significant inequalities between the sexes remain, especially with women still accounting for a disproportionate number of unpaid household workers and informal sector employees (Wacharaporn 2008).

These changes provide the backdrop for the current study, which has examined female labour

[6]Since education category serves as the instrumental variable in the IV estimation, it is a weak instrument in these regressions when the data are disaggregated by education category. There is insufficient variation in the education proportions over time within each education category to effectively identify the labour supply equations. Therefore, IV estimates are not reported for Table 10.

[7]The data in this paragraph are from Nakavachara (2010).

Table 9. The relationship between work and wage for urban/rural women by age.

	Urban			Rural		
Age Group	WLS	IV	E_P	WLS	IV	E_P
25–35	−0.283*	−0.295*	0.127*	−0.266*	−0.299*	0.135*
	(0.0183)	(0.0329)	(0.0161)	(0.0240)	(0.0521)	(0.0175)
36–45	−0.233*	−0.213*	0.0912*	−0.242*	−0.219*	0.0982*
	(0.0102)	(0.0196)	(0.0116)	(0.0102)	(0.0284)	(0.0088)
46–60	−0.182*	−0.116*	0.0876*	−0.165*	−0.127*	0.0887*
	(0.0067)	(0.0338)	(0.0253)	(0.0088)	(0.0537)	(0.0229)
All	−0.225*	−0.218*	0.0834*	−0.218*	−0.226*	0.110*
	(0.0090)	(0.0224)	(0.0109)	(0.0094)	(0.0331)	(0.0100)

Heteroscedastic-consistent estimated standard errors are in parentheses. WLS stands for weighted least-squares.
IV stands for instrumental variable, which is the proportion of women of each cohort in three educational categories.

- Urban – Weak instruments test on first stage regression: Robust F (1,91) = 27.86 (p = 0.00)
  - Age group (25–35) – Weak instruments test on first stage regression: Robust F (1,25) = 34.47 (p = 0.00)
  - Age group (36–45) – Weak instruments test on first stage regression: Robust F (1,28) = 25.10 (p = 0.00)
  - Age group (46–60) – Weak instruments test on first stage regression: Robust F (1,34) = 5.88 (p = 0.02)
- Rural – Weak instruments test on first stage regression: Robust F (1,91) = 24.90 (p = 0.00)
  - Age group (25–35) – Weak instruments test on first stage regression: Robust F (1,25) = 31.42 (p = 0.00)
  - Age group (36–45) – Weak instruments test on first stage regression: Robust F (1,28) = 28.57 (p = 0.00)
  - Age group (46–60) – Weak instruments test on first stage regression: Robust F (1,34) = 3.10 (p = 0.08)

*Significant at 0.05 significant level.

Table 10. Estimates of the association between work and wage by education level.

Education Level	Number of Observations	WLS	E_P
Some primary school and lower	31	−0.183*	−0.0584
		(0.0317)	(0.0632)
Some secondary, high school and lower	31	−0.205*	0.0915
		(0.0246)	(0.0479)
Some university and university degree	31	−0.110*	0.192*
		(0.0150)	(0.0449)
All	93	−0.223*	0.111*
		(0.0806)	(0.0137)

Heteroscedastic-consistent estimated standard errors are in parentheses. WLS stands for weighted least-squares.
*Significant at 0.05 significant level.

supply relations in Thailand estimated over this same 20-year period. Thailand offers an interesting case study of female labour supply relations as a country that has promoted female education and accommodated an expansion of women in the workforce.

One of the most notable findings for female labour supply in developing countries is the inverse relation between hours worked and wages. This empirical regularity is confirmed in this study using methodologies that deal with problems of individual heterogeneity and endogeneity that plague individual data analyses. The estimated wage elasticity in the hours worked equation is robust to a variety of estimation methods and choice of control variables. For the complete sample of women reporting earnings in the Labour Force Surveys, the estimated wage elasticity of hours worked is approximately −0.23. This negative elasticity implies a desire to reduce work hours in favour of other household activities, when an increase in women's wages permits. Estimates of the response of the employment rate are equally robust, implying a 10 per cent increase in female earnings induces a 1 percentage point rise in the employment/population ratio.

This inverse relation between hours worked and wages holds under several alternative disaggregations of the data by age, birth cohort, educational attainment, marital status, or urban-rural residence. In addition, the disaggregated estimations provide further information on female labour supply relations in Thailand. For example, younger female workers tend to be more responsive to wage changes on both intensive and extensive margins of labour supply. This has interesting implications for labour policies such as a minimum wage. Imposing or increasing a minimum wage would tend to increase labour force participation, especially for younger women, while reducing their hours worked. Based on the theory behind the empirical analysis, these are utility increasing responses. The substitution

of household duties and child care for working hours, particularly for young mothers, may have societal benefits as well as advantages to the individual household.

Allowing elasticities to differ by education level and age, younger cohorts of women and those with the highest education show the greatest (positive) response of the employment rate to wage changes. The employment ratios are also highest for those women with the highest levels of education (Table 10). Looking towards the future, continued rising levels of education could increase the supply of these highly productive workers, adding to overall productivity and per capita incomes.

Allowing the labour supply elasticities to differ by urban-rural residence, only minor differences emerge. This is contrary to the conventional view that the early stages of industrialization produce urban jobs that are not favourable to female workers. The finding that women living in urban areas are just as responsive to wage change in their labour supply decisions as are rural residents undermines this hypothesis. In addition, unmarried women show a greater absolute response of hours worked to wage changes, compared with married women. Facing a cut in wages, unmarried women need to shore up their households' income by increasing their hours worked. To the extent that the government provides income support to low wage households, the greatest impact of such assistance would be felt by unmarried women. Tax or subsidy policies could be designed to provide the greatest benefit to single women, to reflect their stronger need for assistance. Married women would also increase their hours worked in this case, but by a lesser amount because of the income cushion provided by their husband's earnings. Contrary to the trailing spouse perspective, which views the married woman to be relatively disadvantaged in the labour market, the estimates in this study indicate that marriage provides protection against adverse wage conditions, allowing married women greater control over the choice between hours of work and other household activities.

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